

**ESSAYS ON THE IDENTIFICATION OF FISCAL
POLICY BEHAVIOR**

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Bing Li

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Abstract: The identification of fiscal authority's stance on debt stabilization has become of crucial practical importance. This dissertation studies two different approaches to identify such fiscal policy behavior. The first chapter shows that the current limited-information approaches to identify such fiscal policy behavior, without taking into account the intertemporal equilibrium condition (IEC) of debt valuation and the monetary policy, are, in general, subject to a simultaneity bias problem and, therefore, the corresponding inferences on such fiscal policy behavior may be misleading. The current approaches used include the Ordinary Least Squares (OLS) regression, the Generalized Method of Moments (GMM) and the bivariate Vector Autoregression (VAR). To correctly identify such fiscal policy behavior, the first chapter proposes a simultaneous equations approach, which incorporates both the IEC and the monetary policy and avoids the simultaneity bias problem a priori. Based on the post-World War II U.S. data, the results show that the U.S. primary surplus' response to the state of government debt is, on average, much weaker than those estimated by the OLS and GMM methods. More importantly, the estimated response is statistically insignificant, which is completely different from the OLS and GMM methods. The results also show that the contemporaneous response of the U.S. total government liabilities to a positive surplus shock is not significantly negative, which is in sharp contrast to the typical results of the bivariate VAR.

In the second chapter, it is recognized that the identification of such fiscal policy behavior is a general equilibrium problem, which requires that monetary policy behavior always be considered simultaneously. Therefore, the second chapter takes a holistic view and jointly identifies monetary and fiscal policy behavior for the U.S. by estimating a standard New-Keynesian sticky-price model with Bayesian methods. By applying Bayesian model comparison techniques to the U.S. Pre-Volcker and Post-1982 samples, the second chapter finds out that (1) both samples favor determinacy over indeterminacy; (2) active fiscal policy is not detected in either sample. The findings are consistent with the previous literature for the Post-1982 sample, but not for the Pre-Volcker sample.



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Chapter 1

On the Identification of Fiscal Policy Behavior

Abstract

The identification of fiscal authority's stance on debt stabilization has become of crucial practical importance. This paper shows that the current limited-information approaches to identify such fiscal policy behavior, without taking into account the intertemporal equilibrium condition (IEC) of debt valuation and the monetary policy, are, in general, subject to a simultaneity bias problem and, therefore, the corresponding inferences on such fiscal policy behavior may be misleading. The current approaches used include the Ordinary Least Squares (OLS) regression, the Generalized Method of Moments (GMM) and the bivariate Vector Autoregression (VAR). To correctly identify such fiscal policy behavior, this paper proposes a simultane-

ous equations approach, which incorporates both the IEC and the monetary policy and avoids the simultaneity bias problem a priori. Based on the post-World War II U.S. data, the results show that the U.S. primary surplus' response to the state of the government debt is, on average, much weaker than those estimated by the OLS and GMM methods. More importantly, the estimated response is statistically insignificant, which is completely different from the OLS and GMM methods. The results also show that the contemporaneous response of the U.S. total government liabilities to a positive surplus shock is not significantly negative, which is in sharp contrast to the typical results of the bivariate VAR.

1.1 Introduction

In the empirical macro literature, simple fiscal policy rules are often used to characterize fiscal policy behavior, just as the Taylor rule being used to characterize monetary policy behavior. In general, a simple fiscal policy rule can be expressed as

$$\tau_t = \gamma_b b_{t-1} + \gamma_Z' Z_t + \psi_t \quad (1.1)$$

in which τ_t is the primary surplus, b_{t-1} is the real value of lagged government debt, Z_t is a vector of other controlled variables, e.g. the output gap, and ψ_t is a fiscal shock.

Recently, the identification of fiscal authority's stance on debt stabilization¹ has

¹This paper only focuses on one aspect of fiscal policy behavior, i.e., the fiscal authority's stance on debt stabilization. The other aspects of fiscal policy behavior, such as the fiscal authority's

become of crucial practical importance. One typical example is that, by identifying such fiscal policy behavior, people are able to draw inferences on the sources of fiscal financing in the U.S. that have kept the U.S. debt to GDP ratio from exploding in the past several decades, as depicted in Figure 1.1. In practice, several ways exist as to how to make the identification, depending on what estimation strategy is used. One typical strategy is to estimate a single-equation fiscal policy rule similar to (1.1) using the Ordinary Least Squares (OLS) regression or the Generalized Method of Moments (GMM) and then draw inferences based on the estimated coefficient of the lagged government debt. Another strategy is to embed a fiscal policy rule similar to (1.1) into a bivariate Vector Autoregression (VAR) and draw inferences based on the impulse response functions².

However, the results of this paper show that methods such as the OLS, GMM and bivariate VAR are, in general, subject to a simultaneity bias problem and, therefore, the corresponding inferences on the fiscal policy behavior may be misleading. To understand the nature of the problem, the key is the intertemporal equilibrium condition (IEC)³ of debt valuation that holds in the general equilibrium of any macro economy regardless of the policy behavior in place. Generally, an IEC is

stance on economic stabilization, are not studied in this paper.

²Since these methods only utilize partial information from the whole economic structure, I label them the limited-information approaches. Typical examples include Bohn (1998) for the OLS, Ballabriga and Martinez-Mongay (2003) and Claeys (2006, 2008) for the GMM and Canzoneri, Cumby and Diba (2001, 2002) and Creel and Bihan (2006) for the bivariate VAR. There is another line of research in the recent literature that utilizes all the information of the economic structure to identify fiscal policy behavior. I label this literature the full-information approach. Typical examples include Caivano (2007), Leeper, Plante and Traum (2009) and Li (2010b), among others. Note, this paper only studies the limited-information approaches.

³This terminology follows Leeper (2009). It is also called the "intertemporal government budget constraint" in the literature, especially by opponents of the fiscal theory of the price level.

expressed as

$$\begin{aligned} \text{Real Market Value of Government Debt} = \\ \text{Expected Present Value of Future Primary Surpluses} + \text{Seigniorage} \end{aligned} \quad (1.2)$$

in which

$$\begin{aligned} \text{Primary Surpluses} = & \text{Total Revenues} \\ & - \text{Government Consumption \& Investment} - \text{Government Transfer Payments} \end{aligned}$$

From the IEC, it is clear that the real value of government debt, which is intrinsically determined by the joint processes of future monetary and fiscal policy behavior, is forward-looking in nature. Therefore, fiscal policy rules similar to (1.1) should always be estimated along with the IEC and the monetary policy rules. However, all of the aforementioned methods fail to do so, which results in the simultaneity bias problem.

The first contribution of this paper is the analytical illustration of the simultaneity bias problem associated with the OLS regression within a simple DSGE model. A few results are worth emphasizing here. First, the simultaneity bias problem associated with the OLS regression is prevailing in the parameter space. Specifically, with passive monetary policy (PM)⁴ and active fiscal policy (AF) in place, the simultaneity bias can be either negative or positive, depending on the particu-

⁴According to Leeper's (1991) terminology, monetary policy is passive when the interest rate responds to inflation at a value of less than one-for-one and active when the response is at a value of more than one-for-one. Fiscal policy is passive when the tax responds to the outstanding government debt strongly and active otherwise. According to Woodford's (1995) terminology, an equilibrium is called non-Ricardian when the monetary policy is passive and the fiscal policy is active. Symmetrically, an equilibrium is called Ricardian when the monetary policy is active and the fiscal policy is passive.

lar parameterization. With active monetary policy (AM) and passive fiscal policy (PF) in place, the simultaneity bias is always negative. Second, the severity of the simultaneity bias problem associated with the OLS regression is related to the underlying monetary policy behavior. In particular, an active fiscal policy is very hard to identify using the OLS regression when the monetary policy is passive⁵.

As a second contribution, the results show that simply adding lags of the primary surplus to a fiscal policy rule as explanatory variables does not necessarily soak up the serial correlation in the fiscal shock, which is sufficient for the simultaneity bias problem. Even when it soaks up the serial correlation, the lagged government debt is still likely to forecast future fiscal shocks and primary surpluses. In either case, the simultaneity bias problem is inevitable. Therefore, modeling fiscal policy inertia by including lagged surpluses in a fiscal policy rule is not a robust solution to the simultaneity bias problem.

To solve the simultaneity bias problem fundamentally, this paper proposes a simultaneous equations approach, which incorporates both the IEC and the monetary policy and forms the third contribution. Following the methodology advocated in Chung and Leeper (2009), the cross-equation restrictions implied by the IEC are imposed on an identified VAR that includes both the monetary and the fiscal policy rules. Based on the post-World War II U.S. data, the results show that the response of the U.S. primary surplus to the state of government debt is, on average, much weaker than those estimated by the OLS and GMM methods. More importantly,

⁵To some extent, this may explain why the empirical literature of the U.S. fiscal policy behavior using the OLS regression (e.g., Bohn (1998)) rarely finds significant evidence favoring the fiscal theory of the price level (i.e., PM/AF).

the estimated response is statistically insignificant, which is completely different from the OLS and GMM methods. The results also show that the contemporaneous response of the U.S. total government liabilities to a positive surplus shock is not significantly negative, which is in sharp contrast to the typical results of the bivariate VAR.

1.2 Related Literature

Depending on different estimation strategy, the current literature of identifying fiscal authority's stance on debt stabilization can be categorized into limited-information approach and full-information approach. This section provides a review of the related literature.

As the simplest limited-information approach, Bohn (1998) estimates a fiscal policy rule as following by the OLS regression based on the U.S. annual data.

$$\tau_t = \rho \cdot b_{t-1} + \alpha_0 + \alpha_G \cdot GVAR_t + \alpha_Y \cdot YVAR_t + \epsilon_t \quad (1.3)$$

in which τ_t is the ratio of primary surplus (taxes minus noninterest spending) to GDP, b_{t-1} is the ratio of (start-of-period) debt to GDP, $GVAR_t$ is the level of temporary government spending and $YVAR_t$ is a business cycle indicator. The benchmark regression in Bohn (1998) is based on the sample from 1916 to 1995. Depending on different subperiods used in the estimation, Bohn (1998) obtains estimates of ρ all between 2.8 and 5.4 percent and argues that the estimates are

all significantly positive and quantitatively reasonable. In conclusion, Bohn (1998) states that the U.S. government has been systematically responding to the state of government debt and so the U.S. fiscal policy is sustainable. As well known, the OLS estimate of ρ is consistent only if b_{t-1} is orthogonal to ϵ_t . This assumption seems to be “justifiable” because b_{t-1} is predetermined relative to ϵ_t . However, this paper illustrates that the orthogonality assumption is not generally valid because the government debt is forward-looking in nature. With serially correlated fiscal shocks, it can be shown that the lagged government debt forecasts future fiscal shocks, causing the simultaneity bias problem and leaving the OLS-based inferences unreliable⁶.

As a generalization of the simple fiscal policy rules similar to (1.1), some people introduce policy inertia to the fiscal policy rules by adding lagged primary surpluses as explanatory variables⁷. This, as they argue, is according to the institutional feature of fiscal policy, i.e. the fiscal policy instruments adjust only gradually towards their targets. It is shown in this paper that a serially correlated fiscal shock is sufficient for the lagged government debt to forecast future fiscal shocks, which causes the simultaneity bias problem. People in this literature usually argue that the inclusion of lagged surpluses is able to soak up the serial correlation in the fiscal shock. Therefore, they consider the fiscal shock as unforecastable and the simultaneity bias problem as avoided. In practice, one typical approach is to estimate a linear fiscal

⁶Similar comments on Bohn (1998)’s OLS regression have been made by Cochrane (1998), Woodford (1998, 2001), Canzoneri, Cumby and Diba (2001) and Davig, Leeper and Chung (2007), among others.

⁷Typical examples include Ballabriga and Martinez-Mongay (2003), Galí and Perotti (2003), Claeys (2006, 2008) and Daniel and Shiamptanis (2008), among others.

policy rule by the instrumental variable (IV) regression, e.g. Galí and Perotti (2003). Instrumental variables are usually used for the contemporaneous output gap, which is considered as endogenous and included to capture the fiscal authority's stance on economic stabilization. The lagged government debt is generally not instrumented because it is assumed to be orthogonal to the fiscal shock. Another typical approach is to estimate a nonlinear fiscal policy rule by the GMM method, e.g. Ballabriga and Martinez-Mongay (2003) and Claey's (2006, 2008). Since the lagged government debt is assumed to be orthogonal to the fiscal shock, it is usually included in the set of instrumental variables. For example, Claey's (2006) estimates the following fiscal policy rule for the G-3 and large European countries.

$$\tau_t = \rho\tau_{t-1} + (1 - \rho) [\delta + \gamma x_t + \beta\pi_t + \omega i_t + \theta b_{t-1}] + \varepsilon_t \quad (1.4)$$

in which τ_t is the primary surplus (net taxes minus noninterest spending), x_t is the output gap, π_t is the inflation rate, i_t is the short-term nominal interest rate and b_{t-1} is the lagged government debt. By treating b_{t-1} as an instrumental variable, Claey's (2006) estimates the fiscal policy rule for the U.S. based on the annual data from 1964 to 2003. The estimate of θ , which is the long-run response of the primary surplus to the lagged government debt, is 0.05 and is significantly positive. In conclusion, Claey's (2006) confirms the qualitative result of Bohn (1998) and argues that U.S. fiscal policy is sustainable or Ricardian. However, this paper illustrates that simply adding lagged surpluses to the fiscal policy rule does not necessarily soak up the serial correlation in the fiscal shock. Even when it soaks up the serial correlation

in the fiscal shock, the lagged government debt is still likely to forecast future fiscal shocks. In either case, the orthogonality condition for the lagged government debt and future fiscal shock is invalid and the simultaneity bias problem is inevitable, which undermines the reliability of the inferences drawn by the IV and the GMM methods.

Other than the OLS, IV and GMM methods, which belong to the single-equation framework, some people in the literature fit a simple fiscal policy rule similar to (1.1) in a bivariate VAR and draw inferences based on the impulse response functions⁸. For example, Canzoneri, Cumby and Diba (2001) estimate and identify a bivariate VAR with the primary surplus to GDP ratio and the total government liabilities (sum the net federal debt and the money base) to GDP ratio based on the U.S. annual data from 1951 to 1995. In their Figure 3 on page 1228, they report that a positive shock to the surplus is followed by a persistent decline in the government liabilities. Since the surplus seems to retire the government liabilities, they conclude that a Ricardian interpretation of the U.S. data is “more plausible” than a non-Ricardian one. However, this paper argues that the simultaneity bias problem found in the single-equation framework may also affect the bivariate VAR, because the latter method involves equation-by-equation regressions.

For the full-information approach, there is a line of research in the recent literature that identifies fiscal policy behavior by directly estimating a full-blown dynamic stochastic general equilibrium (DSGE) model, in which simple fiscal policy rules are

⁸Typical examples include Canzoneri, Cumby and Diba (2001, 2002) and Creel and Bihan (2006), among others.

clearly specified. For example, Caivano (2007), Leeper, Plante and Traum (2009) and Li (2010b) apply the Bayesian methods to estimate DSGE models based on the U.S. data. By this means, they get estimated simple fiscal policy rules. As an advantage, the full-information approach incorporates the IEC and the monetary policy⁹ intrinsically. So it is free of the simultaneity bias problem being examined in this paper. As a caveat, the estimated DSGE model is often too stylized, which implies that the full-information approach is not robust to the potential model misspecification.

To correctly identify fiscal authority's stance on debt stabilization, this paper proposes a simultaneous equations approach by following the methodology advocated in Chung and Leeper (2009). Specifically, the simultaneous equations approach imposes the cross-equation restrictions implied by the IEC on an identified VAR that incorporates both the monetary and the fiscal policy rules and then estimates the fiscal policy rules from the identified VAR. As a limited-information approach, the proposed method is better than the others in the same kind because it takes both the IEC and the monetary policy into account and, therefore, is free of the simultaneity bias problem a priori. On the other hand, it imposes the least amount of economic structure in the estimation. So it is more robust to the potential model misspecification than the full-information approach. To identify the structural VAR, this paper also adopts the identification scheme that is proposed by Blanchard and Perotti (2002) and extended by Perotti (2004). Based on the postwar U.S. data, the simultaneous equations approach delivers results that are qualitatively different

⁹As an exception, Leeper, Plante and Traum (2009) estimates a standard RBC model that does not have monetary policy.

from the previous literature. Therefore, this paper concludes that the simultaneity bias is an essential problem in the identification of fiscal policy behavior.

1.3 OLS Regression and the Simultaneity Bias: An Analytical Investigation

In this section, I use a simple DSGE model similar to Leeper (1991, 2005) to illustrate the simultaneity bias problem associated with the OLS regression. The findings are twofold. First, the existence of simultaneity bias is generic, which can be positive or negative over the parameter space. Second, the severity of simultaneity bias problem is related to the underlying monetary policy behavior. The latter finding advocates identifying fiscal policy jointly with monetary policy, which is supposed to deliver more reliable inferences on fiscal policy behavior.

1.3.1 An Illustrative DSGE Model

Model Setup

This is an endowment economy model, in which an infinitely lived representative agent chooses sequences $\{c_t, M_t, B_t\}_{t=0}^{\infty}$ to solve the household problem:

$$\max E_0 \sum_{t=0}^{\infty} \beta^t [\ln(c_t) + \delta \ln(M_t/P_t)] \quad (1.5)$$

subject to the budget constraint

$$c_t + \frac{M_t + B_t}{P_t} + \tau_t = y + \frac{M_{t-1} + R_{t-1}B_{t-1}}{P_t} \quad (1.6)$$

taking the initial wealth $M_{-1} + R_{-1}B_{-1} > 0$ and sequences $\{y, \tau_t, R_t, P_t\}_{t=0}^{\infty}$ as given.

In (1.5) and (1.6), $\beta \in (0, 1)$ is the discount factor, $\delta \in (0, \infty)$ is the weight on real money balance in the utility function, c_t is the real consumption, M_t is the nominal money balance, B_t is the nominal one-period government debt with gross nominal interest rate R_t , P_t is the price level, y is the endowment that is assumed to be constant and τ_t is the lump-sum taxes (if positive) or transfers (if negative).

There is a government choosing policy sequences $\{M_t, B_t, \tau_t\}_{t=0}^{\infty}$ subject to the government budget constraint

$$\frac{M_t + B_t}{P_t} + \tau_t = g_t + \frac{M_{t-1} + R_{t-1}B_{t-1}}{P_t} \quad (1.7)$$

in which g_t is the real government spending.

The aggregate resource constraint is given by

$$c_t + g_t = y \quad (1.8)$$

For simplicity, I assume $g_t = 0$ for all t . So τ_t represents the primary surplus (if positive) or the primary deficit (if negative). Also, (1.8) reduces to $c_t = c = y$, which is the goods market clearing condition.

The model is closed by specifying simple rules for monetary and fiscal policy behavior, which determine the corresponding policy instruments. Specifically, monetary policy is described by the Taylor-type interest rate rule

$$R_t = e^{\alpha_0} \pi_t^\alpha \theta_t \quad (1.9)$$

in which $\pi_t \equiv P_t/P_{t-1}$ is the inflation rate. Likewise, fiscal policy is described by the tax rule

$$\tau_t = e^{\gamma_0} b_{t-1}^\gamma \psi_t \quad (1.10)$$

in which $b_{t-1} \equiv B_{t-1}/P_{t-1}$ is the real government debt. I assume that the exogenous monetary and fiscal policy shocks θ_t and ψ_t have unit steady state values and their logarithms follow AR(1) processes

$$\ln(\theta_t) = \rho_\theta \ln(\theta_{t-1}) + \varepsilon_t^\theta \quad (1.11)$$

$$\ln(\psi_t) = \rho_\psi \ln(\psi_{t-1}) + \varepsilon_t^\psi \quad (1.12)$$

in which $|\rho_\theta| < 1$ and $|\rho_\psi| < 1$ are assumed for stationarity. The innovations to both policy shocks, ε_t^θ and ε_t^ψ , are assumed to be random variables with zero means and bounded support. Their standard deviations are σ_θ and σ_ψ , respectively. I also assume that ε_t^θ and ε_t^ψ are both serially uncorrelated and uncorrelated between each other for all leads and lags¹⁰.

¹⁰Here I make a strong assumption that ε_t^θ and ε_t^ψ are mutually uncorrelated in order to separate effects of different shocks in a clear manner, which makes the illustration more transparent. To allow for nonzero cross-correlation between different policy shock innovations, even though more realistic, makes the derivation more cumbersome but would not change the main results of this paper significantly.

Determinacy Regions

With the complete model setup, I log-linearize the equilibrium conditions around the stationary steady state and then solve the linearized model¹¹. As shown in Appendix 1.A.1, the determinacy of bounded equilibrium¹² of the linearized system hinges on (α, γ) , the two parameters in the monetary and fiscal policy rules. Following Leeper (1991, 2005), I characterize four different regions in the first quadrant of (α, γ) space¹³, two of which indicate determinacy: (1) When $\alpha > 1$ (AM) and $\gamma > 1$ (PF), the equilibrium is determinate; (2) When $\alpha < 1$ (PM) and $\gamma < 1$ (AF), the equilibrium is also determinate; (3) When $\alpha < 1$ (PM) and $\gamma > 1$ (PF), bounded equilibrium is indeterminate; (4) When $\alpha > 1$ (AM) and $\gamma < 1$ (AF), bounded equilibrium does not exist. For the rest of the paper, I only focus on the determinacy regions, which are displayed in Figure 1.2.

1.3.2 OLS Regression and the Simultaneity Bias

For illustration, the DSGE model specified earlier is taken as the data-generating process (DGP). Suppose an econometrician, as in Bohn (1998), tries to identify fiscal authority's stance on debt stabilization by running OLS regression on the simple

¹¹See Appendix 1.A.1 for full description of the model solution.

¹²In this paper, I only focus on determinate bounded equilibrium. The earlier assumption that random variables ε_t^θ and ε_t^ψ have bounded support forms a necessary condition for this.

¹³As to the determinacy regions in the (α, γ) space, I only focus on those in the first quadrant, which are economically more meaningful than those in the other quadrants.

fiscal policy rule, which, in log-linearized form, is given by

$$\hat{\tau}_t = \gamma \hat{b}_{t-1} + \hat{\psi}_t \quad (1.13)$$

in which \hat{x}_t denotes log deviation of the generic variable x_t from its steady state value x , i.e. $\hat{x}_t \equiv \ln(x_t) - \ln(x)$. To simplify the analysis, the econometrician is assumed to know the correct functional form of the fiscal policy rule (1.13). From the econometrician's perspective, \hat{b}_{t-1} is predetermined relative to $\hat{\psi}_t$. Therefore, \hat{b}_{t-1} is considered as orthogonal to $\hat{\psi}_t$ and the OLS estimator of γ , i.e. $\hat{\gamma}_{OLS}$, is considered as "consistent". Since (1.13) is a structural equation, the econometrician naturally attaches behavioral interpretation to $\hat{\gamma}_{OLS}$ and thus identifies the underlying fiscal policy behavior.

Mathematically, the probability limit of $\hat{\gamma}_{OLS}$ is

$$\text{plim} \hat{\gamma}_{OLS} = \gamma + \frac{\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t)}{\text{var}(\hat{b}_{t-1})} \quad (1.14)$$

in which $\text{cov}(\cdot, \cdot)$ and $\text{var}(\cdot)$ are the covariance and variance operators, respectively. It is clear from (1.14) that $\hat{\gamma}_{OLS}$ is consistent only if the term $\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t)/\text{var}(\hat{b}_{t-1})$, which is the slope coefficient in the regression of $\hat{\psi}_t$ on \hat{b}_{t-1} , is zero. In another word, $\hat{\gamma}_{OLS}$ is inconsistent if past government debt forecasts future fiscal shocks. In this section, I illustrate that past government debt does forecast future fiscal shocks over a wide range of the parameter space of the illustrative model, with exceptions only in some rare cases. This implies that $\hat{\gamma}_{OLS}$ is generally inconsistent and the

corresponding identification of fiscal policy behavior may be unreliable. The key to understand the subtle point lies in the fact that \hat{b}_{t-1} is forward-looking in nature, which is best demonstrated by the IEC that can be obtained by iterating (2.27) forwards over B/P , taking expectation conditional on the information set at $t - 1$ and imposing the transversality condition from the household problem:

$$\frac{B_{t-1}}{P_{t-1}} = E_{t-1} \sum_{j=0}^{\infty} \left(\prod_{k=0}^j \pi_{t+k} R_{t+k-1}^{-1} \right) \left[\tau_{t+j} + \frac{M_{t+j} - M_{t+j-1}}{P_{t+j}} \right] \quad (1.15)$$

The IEC (1.15) is a general equilibrium condition that always holds in equilibrium regardless of the policy behavior in place. It basically says that, in equilibrium, the real market value of government debt is always equal to the expected present value of future primary surpluses, inclusive of seigniorage revenues. When past government debt forecasts future fiscal shocks, the correlation between \hat{b}_{t-1} and $\hat{\tau}_t$ not only reflects the dependence of $\hat{\tau}_t$ on \hat{b}_{t-1} through γ of the fiscal policy rule but also carries the correlation between \hat{b}_{t-1} and $\hat{\psi}_t$ that is captured by the IEC. In equilibrium, these two parts of the correlation between \hat{b}_{t-1} and $\hat{\tau}_t$ hold in the fiscal policy rule and the IEC simultaneously. Hence, if the econometrician isolates the fiscal policy rule from the whole system of structural equations, particularly the IEC, and projects $\hat{\tau}_t$ on \hat{b}_{t-1} , the econometric model is misspecified and the OLS estimator of γ is subject to simultaneity bias.

Next, I illustrate the nature of the simultaneity bias problem through four simple cases. The DGPs of the first three cases are characterized by PM/AF policy combination and that of the last case is characterized by AM/PF.

Case I (PM/AF): $\alpha = \gamma = \rho_\theta = \rho_\psi = 0$

In this case, the nominal interest rate is pegged, the primary surplus is exogenous and both the monetary and the fiscal shocks are serially uncorrelated. Under these simplifying assumptions, I get

$$\hat{b}_t = \beta(\varphi_1 + \varphi_4)\varepsilon_t^\theta \quad (1.16)$$

in which, according to Appendix 1.A.1, $\varphi_1 \equiv m(\alpha\chi + 1)/b + \beta^{-1}$, $\varphi_4 \equiv -m\chi/b - \beta^{-1}$ and $\chi \equiv 1/(1 - R)$. Since $\rho_\psi = 0$, I get $\hat{\psi}_t = \varepsilon_t^\psi$. Hence, the covariance term in (1.14) can be shown as

$$\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t) = \text{cov}\left[\beta(\varphi_1 + \varphi_4)\varepsilon_{t-1}^\theta, \varepsilon_t^\psi\right] = 0 \quad (1.17)$$

The covariance is zero because of the assumption that ε_t^θ and ε_t^ψ are both serially uncorrelated and uncorrelated between each other for all leads and lags. Putting in another way, the only shock that impinges on \hat{b}_{t-1} is the monetary shock ε_{t-1}^θ , which is by assumption orthogonal to future fiscal shocks. Therefore, \hat{b}_{t-1} is orthogonal to $\hat{\psi}_t$, i.e. government debt does not forecast future fiscal shocks. In consequence, $\text{plim}\hat{\gamma}_{OLS} = \gamma$ and the OLS estimator of γ is free of simultaneity bias.

However, the assumption that $\hat{\psi}_t$ in the simple rule (1.13) has no serial correlation is too restrictive and is not consistent with the existing empirical evidence. Therefore, more realistic cases are explored as follows.

Case II (PM/AF): $\alpha = \gamma = \rho_\theta = 0$, $0 < \rho_\psi < 1$

In this case, I keep the assumptions of case I except that fiscal policy shock $\hat{\psi}_t$ becomes serially correlated. Under these assumptions, I get

$$\begin{aligned}\hat{b}_t &= \beta(\varphi_1 + \varphi_4)\hat{\theta}_t + \left[\frac{(\beta^{-1} - 1)\rho_\psi}{\beta^{-1} - \rho_\psi} \right] \hat{\psi}_t \\ &= \beta(\varphi_1 + \varphi_4)\varepsilon_t^\theta + \left[\frac{(\beta^{-1} - 1)\rho_\psi}{\beta^{-1} - \rho_\psi} \right] \frac{\varepsilon_t^\psi}{(1 - \rho_\psi L)}\end{aligned}\quad (1.18)$$

in which $\hat{\psi}_t \equiv \varepsilon_t^\psi / (1 - \rho_\psi L)$ and L is the lag operator. The covariance term in (1.14) is then given by

$$\begin{aligned}\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t) &= \text{cov} \left[\beta(\varphi_1 + \varphi_4)\varepsilon_{t-1}^\theta + \frac{(\beta^{-1} - 1)\rho_\psi}{\beta^{-1} - \rho_\psi} \frac{\varepsilon_{t-1}^\psi}{(1 - \rho_\psi L)}, \frac{\varepsilon_t^\psi}{(1 - \rho_\psi L)} \right] \\ &= \left[\frac{(\beta^{-1} - 1)\rho_\psi}{\beta^{-1} - \rho_\psi} \right] \text{cov} \left[\frac{\varepsilon_{t-1}^\psi}{(1 - \rho_\psi L)}, \frac{\varepsilon_t^\psi}{(1 - \rho_\psi L)} \right] \\ &= \left[\frac{(\beta^{-1} - 1)\sigma_\psi^2}{\beta^{-1} - \rho_\psi} \right] \sum_{i=1}^{\infty} \rho_\psi^{2i} \\ &= \frac{(\beta^{-1} - 1)\rho_\psi^2 \sigma_\psi^2}{(\beta^{-1} - \rho_\psi)(1 - \rho_\psi^2)} > 0\end{aligned}\quad (1.19)$$

It is obviously seen that the simultaneity bias would always be positive because of the positive covariance, i.e. $\text{plim} \hat{\gamma}_{OLS} > \gamma$. Hence, the OLS estimator of γ is inconsistent. Even though the fiscal policy is assumed to be active ($\gamma = 0$), asymptotically, $\hat{\gamma}_{OLS}$ may be significantly large under certain parameterizations so that the econometrician may mistakenly identify the underlying fiscal policy as passive.

To get more economic insight into the positive bias, suppose there is a surprise

tax cut or negative surplus shock in period $t - 1$, i.e. $\varepsilon_{t-1}^\psi < 0$. Since $0 < \rho_\psi < 1$ and $\gamma = 0$, a negative surplus shock at $t - 1$ forecasts lower path of $\hat{\psi}_{t+j}$ or $\hat{\tau}_{t+j}$ ($j \geq 0$). Because the ex-ante real interest rate is constant in this economy with fixed endowment, the expected discount factor is constant. Hence, the expected present value of future surpluses decreases. I label it the expected surplus effect. On the other hand, the lower path of expected future taxes immediately raises demand for goods via the wealth effect. With the constant supply of goods, inflation increases on impact. However, the nominal interest rate is pegged and does not respond to higher inflation, which in turn does not raise the expected inflation according to the Fisher equation. So the expected present value of future seigniorage is unaffected. In summary, \hat{b}_{t-1} decreases due to the expected surplus effect. In this case, \hat{b}_{t-1} forecasts smaller $\hat{\psi}_t$ and proves to be positively correlated with $\hat{\psi}_t$, which results in the positive simultaneity bias.

Case III (PM/AF): $\gamma = \rho_\theta = 0$, $0 < \alpha < 1$, $0 < \rho_\psi < 1$

In this case, I keep the assumptions of case II except that the monetary authority starts to respond to inflation when setting the nominal interest rate. The derivation of $\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t)$ and $\text{var}(\hat{b}_{t-1})$ is shown in Appendix 1.A.2. From the algebraic derivation, it is not obvious to determine the sign of the simultaneity bias. For illustration purpose, I instead evaluate the bias numerically over a grid in the (α, ρ_ψ) space, in which both parameters range from 0.01 to 0.99 with increments of 0.01. I normalize the economy by setting $y = 1$. As assumed, γ and ρ_θ are both zeroes. The other

parameters are calibrated to match the U.S. data¹⁴. First, β and δ are calibrated as 0.99 and 0.0007, respectively, to match the steady state quarterly nominal interest rate 1.01 and the real money balance to GDP ratio 0.07. Second, γ_0 is implicitly calibrated to match the steady state real government debt to GDP ratio 0.34. Finally, σ_θ and σ_ψ are calibrated from a Bayesian perspective. Specifically, I construct the series of ε_t^θ and ε_t^ψ at each grid point based on (1.69), (1.74) and (1.77) in Appendix 1.A.1. At each grid point, I then get a unique pair of σ_θ and σ_ψ that is consistent with the U.S. data.

The left panel of Figure 1.3 displays the simultaneity bias over the grid, which is calculated based on the analytical result in Appendix 1.A.2. For clarity, I provide the associated contour plot in the right panel. Conditioning on the calibrated parameter values, it is clear that the bias can be positive or negative¹⁵, depending on α and ρ_ψ . Obviously, with highly persistent fiscal policy shocks (large ρ_ψ) and relatively passive monetary policy (small α), the bias is positive. By keeping ρ_ψ fixed and increasing α , the bias becomes negative. Note, the negative bias is nonlinear in α , which is explained later.

For the economic intuition, I carry out the same thought experiment as in the previous case. Again, suppose there is a surprise tax cut or negative surplus shock in period $t - 1$, which forecasts lower path of future surpluses. This lowers the expected present value of future surpluses on the RHS of (1.15), which is the expected surplus

¹⁴I use the quarterly U.S. data from 1947:2 to 2006:2. See Appendix 1.C.4 for data construction.

¹⁵The purpose of this case is to illustrate that the bias can be positive or negative. The magnitude of the bias is of less importance, because it depends on the particular model setup and the calibrated parameter values. The empirical significance of the simultaneity bias problem will be explored in more depth in section 1.6.

effect. Again, lower path of expected future surpluses leads to higher inflation in period $t - 1$ via the wealth effect. Due to the assumption that $0 < \alpha < 1$, higher inflation leads to higher nominal interest rate through the monetary policy rule. According to the Fisher equation, expected inflation increases, which in turn raises the expected present value of future seigniorage. I label it the expected seigniorage effect, which operates in the opposite direction against the expected surplus effect. In summary, the net effect of a negative surplus shock on \hat{b}_{t-1} depends on the tradeoff between the two effects and, therefore, the correlation between \hat{b}_{t-1} and $\hat{\psi}_t$ is ambiguous. Loosely speaking, given α , higher ρ_ψ indicates stronger expected surplus effect, which tends to reduce \hat{b}_{t-1} . Since \hat{b}_{t-1} forecasts smaller $\hat{\psi}_t$, \hat{b}_{t-1} and $\hat{\psi}_t$ are positively correlated, which results in a positive simultaneity bias. On the other hand, given ρ_ψ , higher α implies stronger expected seigniorage effect, which tends to raise \hat{b}_{t-1} . Since \hat{b}_{t-1} forecasts smaller $\hat{\psi}_t$, \hat{b}_{t-1} and $\hat{\psi}_t$ are negatively correlated, which results in a negative simultaneity bias. These possibilities are displayed in Figure 1.3.

With the above economic interpretation, the nonlinearity of the negative bias as one increases α can be explained by the fact that the expected seigniorage effect is nonlinear in α , which can be further explained by the underlying Laffer curve. As the economy stays on the upward-sloping side of the Laffer curve, the expected seigniorage effect becomes more dominant as α gets larger, which turns the correlation between \hat{b}_{t-1} and $\hat{\psi}_t$ from positive to negative. When the economy gets to the downward-sloping side of the Laffer curve, higher α induces weaker expected seigniorage effect, which makes the expected surplus effect relatively stronger and

drives the bias up towards zero.

Case IV (AM/PF): $\alpha > 1$, $\gamma > 1$, $0 < \rho_\psi < 1$

In this case, monetary policy is active and fiscal policy is passive. Under these assumptions, I get¹⁶

$$\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t) = -\frac{\rho_\psi \sigma_\psi^2}{[1 - (\beta^{-1} - \gamma(\beta^{-1} - 1))\rho_\psi](1 - \rho_\psi^2)} \quad (1.20)$$

which is always negative because $\beta^{-1} - \gamma(\beta^{-1} - 1) < 1$ and $0 < \rho_\psi < 1$. Hence, the simultaneity bias is negative, i.e. $\text{plim} \hat{\gamma}_{OLS} < \gamma$, and the OLS estimator of γ is inconsistent. Symmetrically to case II, $\hat{\gamma}_{OLS}$ may be significantly smaller than γ so that the econometrician may mistakenly identify the underlying fiscal policy as active.

The economic interpretation for this case is straightforward. Again, suppose there is a surprise tax cut or negative surplus shock in period $t - 1$. It can be shown that \hat{b}_{t-1} increases on impact to finance the tax cut. According to the fiscal policy rule, higher \hat{b}_{t-1} raises lump-sum tax in period t , which neutralizes the wealth effect of the tax cut to the point that inflation in period $t - 1$ keeps unaffected. This is the standard story of Ricardian equivalence. As $\rho_\psi > 0$, the increased \hat{b}_{t-1} forecasts a smaller $\hat{\psi}_t$. Therefore, \hat{b}_{t-1} is negatively correlated with $\hat{\psi}_t$, which results in the negative simultaneity bias.

¹⁶See Appendix 1.A.2 for more details.

Summary

From these simple cases, it is clear that the simultaneity bias problem associated with the OLS regression is prevailing in the parameter space as long as \hat{b}_{t-1} forecasts $\hat{\psi}_t$, whatever the underlying DGP is. In fact, this problem could make correct identification of fiscal policy behavior very difficult.

With PM/AF in the DGP, on one hand, fiscal policy could be very active, e.g. γ could be zero. Thus \hat{b}_{t-1} and $\hat{\tau}_t$ have no link through the fiscal policy rule. On the other hand, the covariance between \hat{b}_{t-1} and $\hat{\psi}_t$ could be positive or negative as long as \hat{b}_{t-1} forecasts $\hat{\psi}_t$. This makes the correlation between \hat{b}_{t-1} and $\hat{\tau}_t$ significantly nonzero, which is captured by the IEC. When the econometrician runs OLS regression using equilibrium data, the resulting $\hat{\gamma}_{OLS}$ may be significantly positive or negative, which certainly does not reflect the true fiscal policy behavior. With AM/PF in place, on one hand, the fiscal policy rule implies a positive mapping from \hat{b}_{t-1} to $\hat{\tau}_t$ through γ . On the other hand, the negative covariance between \hat{b}_{t-1} and $\hat{\psi}_t$ could be very strong so that the overall correlation between \hat{b}_{t-1} and $\hat{\tau}_t$ is negative. Overall, $\hat{\gamma}_{OLS}$ messes up the true fiscal policy behavior with the correlation between \hat{b}_{t-1} and $\hat{\tau}_t$ that is captured by the IEC.

1.3.3 Monetary Policy Behavior and the Simultaneity Bias

In this section, I further illustrate the simultaneity bias problem by taking monetary and fiscal policy interaction into account. I find out that the severity of simultaneity

bias problem is related to the underlying monetary policy behavior, especially when fiscal policy is active. An important empirical implication goes as follows. By applying the OLS regression as in Bohn (1998), it is difficult to correctly identify a world consistent with the fiscal theory of the price level (PM/AF). Hence, I argue that correctly identifying monetary policy behavior provides useful information for identifying fiscal policy behavior. Besides, from the perspective of the IEC (1.15), the real value of the government debt is related to monetary policy behavior through the discount factor. Therefore, missing monetary policy in the identification of fiscal policy behavior is inherently subject to the simultaneity bias problem, which confirms the necessity of jointly identifying both the monetary and the fiscal policy behavior.

For illustration, I set up grids in the (α, γ) space, over which I numerically evaluate the simultaneity bias. Specifically, for the region corresponding to PM/AF, both parameters range from 0.01 to 0.99 with increments of 0.01; for the region corresponding to AM/PF, both parameters range from 1.01 to 1.99 with increments of 0.01. For both regions, I calibrate $(y, \beta, \delta, \gamma_0, \rho_\theta, \rho_\psi, \sigma_\theta, \sigma_\psi)$ in the same way as in case III of the previous section. To calibrate $(\rho_\theta, \rho_\psi, \sigma_\theta, \sigma_\psi)$, I construct the series of $\hat{\theta}_t$ and $\hat{\psi}_t$ based on (1.69) and (1.74) in Appendix 1.A.1. I then obtain $(\rho_\theta, \rho_\psi, \sigma_\theta, \sigma_\psi)$ by running simple regressions on (1.76) and (1.77) in Appendix 1.A.1. Thus, for each grid point, I have a unique set of $(\rho_\theta, \rho_\psi, \sigma_\theta, \sigma_\psi)$ that is consistent with the U.S. data.

Based on the analytical result in Appendix 1.A.2, I calculate the simultaneity

bias for both regions, which are displayed in the left panels of Figure 1.4 and 1.5. For clarity, I provide the associated contour plots in the right panels. In Figure 1.4, which is for the region of PM/AF, it is noticeable that the bias is positive when α is relatively small ($\alpha < 0.21$), i.e. when monetary policy is very passive. This is because when α is very small, the expected surplus effect dominates the expected seigniorage effect, which implies positive correlation between \hat{b}_{t-1} and $\hat{\psi}_t$ and thus positive bias. In this area, monetary policy is very passive and the nominal interest rate is nearly pegged. Hence, the volatility of the expected inflation and the expected seigniorage is relatively small compared to the cases in which α is large. From (1.15), it is easy to see that the volatility of \hat{b}_{t-1} is relatively small. According to (1.14), the simultaneity bias is relatively large in magnitude. By fixing γ and increasing α , i.e. monetary policy gets more active, the bias gradually becomes negative because the expected seigniorage effect becomes dominant over the expected surplus effect. However, the negative bias is nonlinear in α , which is indicated by the U-shaped plane. This can be explained by the underlying Laffer curve, which has been mentioned earlier. In figure 1.5, which is for the region of AM/PF, the bias is always negative and turns out to be a downward-sloping plane in γ . This is because as γ gets larger, tax responds to government debt more strongly. This makes the expected seigniorage less volatile because the wealth effect is weaker. According to (1.15), $\text{var}(\hat{b}_{t-1})$ gets smaller and the bias gets larger in magnitude.

From the pattern of the simultaneity bias in Figure 1.4, it is clear that an active fiscal policy is very hard to identify by the OLS regression when the monetary policy in place is very passive. This is because when monetary policy is very passive, such

as interest rate pegging ($\alpha \approx 0$), the OLS estimator of γ is significantly biased up. Considering the empirical literature trying to identify fiscal policy behavior or test fiscal sustainability by the OLS regression, such as Bohn (1998) for the U.S. data, most papers get results favoring passive fiscal policy. In another word, a world consistent with the fiscal theory of the price level (PM/AF) is rarely identified in practice by the OLS regression. Connecting with my analysis, it might be the case that the fiscal policy in place was actually active, which is just not identifiable by the OLS regression because the monetary policy in place was very passive and the OLS estimator of fiscal policy rule is significantly biased up.

1.4 Fiscal Policy Inertia and GMM: An Analytical Investigation

Unlike the OLS regression as in Bohn (1998), another strand of the empirical literature that estimates simple fiscal policy rules proposes to take policy inertia into account. As people in this literature argue, policy inertia is supported by the institutional feature of fiscal policy, i.e. fiscal policy instruments adjust only gradually towards their targets. In practice, people add lagged primary surpluses to the right-hand side of the simple fiscal policy rules as explanatory variables¹⁷. In the context of the illustrative DSGE model specified in the previous section, a typical fiscal

¹⁷In the literature of estimating simple fiscal policy rules, it is very common to add the first lag of the primary surplus as an explanatory variable. Note, the simultaneous equations approach proposed in this paper allows for more general fiscal policy rules that include more lags of the primary surplus.

policy rule with policy inertia is in the following form

$$\hat{\tau}_t = (1 - \omega)\hat{\tau}_{t-1} + \omega\gamma\hat{b}_{t-1} + \hat{\psi}_t \quad (1.21)$$

in which the fiscal authority is allowed to smooth primary surplus by responding to its first lag. $(1 - \omega)$ represents the persistence in the surplus. $\omega\gamma$ and γ represent the short-run and the long-run responses of surplus to lagged government debt, respectively.

Recalling the OLS regression analyzed in the previous section, serial correlation in the fiscal shock is sufficient for \hat{b}_{t-1} to forecast $\hat{\psi}_t$, which results in the simultaneity bias problem. In the literature with built-in policy inertia, people usually argue that $\hat{\tau}_{t-1}$ soaks up the serial correlation in $\hat{\psi}_t$ so that the simultaneity bias problem can be avoided. Therefore, \hat{b}_{t-1} is always assumed to be orthogonal to $\hat{\psi}_t$ implicitly. The related work includes the linear IV estimation as in Galí and Perotti (2003) and the GMM estimation as in Ballabriga and Martinez-Mongay (2003) and Claeys (2006, 2008).

In this section, I illustrate that simply adding $\hat{\tau}_{t-1}$ to the fiscal policy rule does not necessarily soak up the serial correlation in $\hat{\psi}_t$. Even when it soaks up the serial correlation in $\hat{\psi}_t$, \hat{b}_{t-1} is still likely to forecast $\hat{\psi}_t$. Due to the forward-looking nature of government debt, \hat{b}_{t-1} and $\hat{\psi}_t$ are very likely to be correlated, even with policy inertia. This undermines the reliability of the inferences drawn by the IV and the GMM estimation.

For illustration, an example is given by modifying the model specified earlier. The modification is in two aspects. First, the original fiscal policy rule (1.13) is replaced by (1.21). Second, an information assumption on $\hat{\psi}_t$ is made following the literature of fiscal foresight (or anticipated shock)¹⁸, which argues that fiscal shocks may be learnt by economic agents before they fully materialize. A general reason for this information assumption is the legislative lag and the implementation lag that are inherent in the fiscal policy making process. Specifically, I assume the fiscal shock process as

$$\hat{\psi}_t = \rho_\psi \hat{\psi}_{t-1} + \phi \varepsilon_{t-1}^\psi + (1 - \phi) \varepsilon_t^\psi \quad (1.22)$$

in which one-period foresight is assumed¹⁹. The conventional information assumption sets $\phi = 0$. When $\phi = 1$, economic agents have complete one-period foresight, i.e. fiscal shock (news) occurring at period $t - 1$ does not materialize until period t . To simplify the illustration, I also assume $\alpha = \gamma = \rho_\theta = \rho_\psi = 0$, which implies that passive monetary policy and active fiscal policy are in place. So the bounded equilibrium is determinate.

Following the solution procedure in Appendix 1.B, I can show that

$$\hat{b}_t = \frac{(1 - \omega)(1 - \beta)}{1 - \beta + \omega\beta} \hat{\tau}_t + \beta(\varphi_1 + \varphi_4) \varepsilon_t^\theta + \frac{(1 - \beta)\phi}{1 - \beta + \omega\beta} \varepsilon_t^\psi \quad (1.23)$$

¹⁸Examples include Leeper (1989), Yang (2005), Leeper, Walker and Yang (2008, 2009), Schmitt-Grohé and Uribe (2008), among others.

¹⁹One-period foresight is assumed to simplify the illustration. One can certainly generalize (1.22) by specifying a longer foresight, which will not change the nature of the problem examined here.

in which

$$\hat{\tau}_t = (1 - \omega)\hat{\tau}_{t-1} + \hat{\psi}_t = (1 - \omega)\hat{\tau}_{t-1} + \phi\varepsilon_{t-1}^\psi + (1 - \phi)\varepsilon_t^\psi \quad (1.24)$$

Then it is easy to show that

$$\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t) = \left[\frac{(1 - \omega)(1 - \beta)(1 - \phi)\phi + (1 - \beta)\phi^2}{1 - \beta + \omega\beta} \right] \sigma_\psi^2 > 0 \quad (1.25)$$

The economic interpretation of the positive covariance in (1.25) is straightforward. Suppose there is a surprise tax cut or negative surplus shock in period $t - 1$, i.e. $\varepsilon_{t-1}^\psi < 0$ or $\hat{\psi}_{t-1} < 0$. On impact, $\hat{\tau}_{t-1}$ decreases according to (1.24). Although $\hat{\psi}_t$ has no autoregressive term by assumption ($\rho_\psi = 0$), $\hat{\psi}_{t-1}$ and $\hat{\psi}_t$ are still correlated due to the moving average term in (1.22)²⁰. Hence, $\hat{\psi}_t$ is expected to decrease. Because of the decrease in $\hat{\tau}_{t-1}$ and the expected decrease in $\hat{\psi}_t$, $\hat{\tau}_t$ is expected to decrease according to (1.24), which induces the expected surplus effect and reduces the right-hand side of the IEC (1.15). As a result, \hat{b}_{t-1} decreases. In summary, \hat{b}_{t-1} forecasts $\hat{\psi}_t$ and proves to be positively correlated with $\hat{\psi}_t$.

When $\phi = 1$, (1.24) and (1.25) are simplified as

$$\hat{\tau}_t = (1 - \omega)\hat{\tau}_{t-1} + \hat{\psi}_t = (1 - \omega)\hat{\tau}_{t-1} + \phi\varepsilon_{t-1}^\psi \quad (1.26)$$

²⁰With one-period foresight, serial correlation in $\hat{\psi}_t$ only lasts for one period before it disappears, e.g. $\hat{\psi}_{t-1}$ and $\hat{\psi}_t$ are correlated. By introducing higher order fiscal foresight or adding more moving average terms to the fiscal shock process, serial correlation in $\hat{\psi}_t$ can last for longer periods. Note, this generalization will not change the nature of the problem examined here.

and

$$\text{cov}(\hat{b}_{t-1}, \hat{\psi}_t) = \frac{(1 - \beta)\sigma_\psi^2}{1 - \beta + \omega\beta} > 0 \quad (1.27)$$

Obviously, even when $\hat{\psi}_t$ is serially uncorrelated, \hat{b}_{t-1} is still likely to forecast $\hat{\psi}_t$, which results in the simultaneity bias problem.

The central message from this example is that including lagged surpluses in a fiscal policy rule still misses the IEC and the monetary policy behavior and, therefore, is not a robust solution to the simultaneity bias problem. This implies that the empirical results from the IV and the GMM estimation may be unreliable. Taking the GMM estimation as example, with one-period foresight, not only \hat{b}_{t-1} but also other variables such as $\hat{\tau}_{t-1}$ is correlated with $\hat{\psi}_t$. With more moving average terms in the fiscal shock process, more lags of the variables become correlated with $\hat{\psi}_t$. These pitfalls leave a wide range of orthogonality conditions in the GMM estimation questionable.

1.5 Empirical Analysis: The Existing Approaches

To contrast the existing limited-information approaches to the simultaneous equations approach that is introduced later, I first run an OLS regression on a simple fiscal policy rule using the U.S. annual data. Then I apply GMM estimation to a simple nonlinear fiscal policy rule using the quarterly U.S. data. In addition, I use the quarterly U.S. data to estimate and identify a bivariate VAR composed of surplus/GDP and total government liabilities/GDP. The qualitative results from these

estimation are consistent with the main findings of the existing literature.

1.5.1 The OLS Regression

First, I run an OLS regression on a simple fiscal policy rule using the U.S. annual data from 1948 to 1990. The fiscal policy rule is the one proposed by Bohn (1998), as shown in (1.3). For readers' convenience, it is reported as below.

$$\tau_t = \rho \cdot b_{t-1} + \alpha_0 + \alpha_G \cdot GVAR_t + \alpha_Y \cdot YVAR_t + \epsilon_t$$

The data are taken from the technical appendix to Bohn (1998).

As a result, the point estimate of ρ , i.e. $\hat{\rho}_{OLS}$, is 0.041. The 90% confidence interval is $[0.017, 0.065]$, which is computed with Newey-West heteroskedasticity- and autocorrelation- consistent standard error with lag window of size 1. Obviously, the coefficient on lagged government debt is significantly positive and is also quantitatively reasonable according to the argument of Bohn (1998). The value of $\hat{\rho}_{OLS}$ means that a marginal increase in government debt by \$100 would increase the primary surplus in the following year by \$4.10. The qualitative result here is consistent with the main conclusion of Bohn (1998), even though the time span I choose is different from Bohn (1998). Specifically, I may conclude that the U.S. government has been systematically responding to the state of government debt and, therefore, the postwar U.S. fiscal policy is sustainable. According to Woodford (1995)'s terminology, I may infer that the postwar U.S. fiscal policy on average can

be identified as Ricardian.

1.5.2 The GMM Method

Now I apply GMM method to a simple nonlinear fiscal policy rule using the quarterly U.S. data. The fiscal policy rule is the one specified in Claeys (2006), as shown in (1.4). For readers' convenience, it is reported as below.

$$\tau_t = \rho\tau_{t-1} + (1 - \rho) [\delta + \gamma x_t + \beta\pi_t + \omega i_t + \theta b_{t-1}] + \varepsilon_t$$

As in Claeys (2006), all of the real quantity variables $[\tau_t, x_t, b_{t-1}]$ are scaled by the real potential output²¹. In particular, τ_{t-1} enters to capture policy inertia, which is measured by the persistence parameter ρ . The parameter of more interest is θ , which measures the long-run response of primary surplus to lagged government debt and characterizes fiscal authority's stance on debt stabilization.

To apply GMM estimation to the fiscal policy rules like (1.4), the literature generally assumes that τ_{t-1} soaks up potential serial correlation in the fiscal shock so that ε_t is “orthogonal” to the past information set, including the lagged government debt b_{t-j} (for $j \geq 1$). Therefore, a set of “orthogonalization” conditions as (1.28) can be obtained.

$$\{[\tau_t - \rho\tau_{t-1} - (1 - \rho)(\delta + \gamma x_t + \beta\pi_t + \omega i_t + \theta b_{t-1})]\tilde{z}_t\} = 0 \quad (1.28)$$

²¹The real potential output is constructed by the Congressional Budget Office and is downloaded from <http://research.stlouisfed.org/fred2/series/GDPPOT?cid=106>.

in which \tilde{z}_t denotes a set of instrumental variables that are “orthogonal” to ε_t . Specifically, I choose \tilde{z}_t to include $[1, \tau_{t-1}, x_{t-1}, \pi_{t-1}, i_{t-1}, b_{t-1}]$. In addition, I follow Claeys (2006) to include some other variables in \tilde{z}_t , which represent domestic and international monetary conditions and supply-side factors. They are the first lags of the M3, the producer price index (PPI), the 10-year treasury bond yield (i^{Long}), Germany short-term interest rate ($i^{Germany}$), the exchange rate between Germany and the U.S. (DEMUSD) and the NAIRU²².

The baseline GMM estimate of the fiscal policy rule (1.4) is shown in Table 1.1²³. Since the focus is exclusively on θ , only the result of $\hat{\theta}_{GMM}$ is reported. Due to the difference in data availability over different time spans, four cases are investigated based on different sets of instrumental variables. Time span for each case is listed in the first column of Table 1.1. From the baseline result, it is clear that $\hat{\theta}_{GMM}$ is significantly positive and is quantitatively comparable across all the cases. At 5% level, the orthogonalization condition (1.28) is not rejected by Hansen’s J test for all the cases. The qualitative result confirms the argument in Claeys (2006) that U.S. fiscal policy is sustainable or Ricardian.

To check robustness of the baseline result, I implement two experiments. First, I expand the set of instrumental variables for all the cases by including the second lags of the variables. Results are shown in Table 1.2. Obviously, this experiment does not change the baseline result either qualitatively or quantitatively. Second, I estimate

²²The M3, PPI and DEMUSD are from Federal Reserve Bank of St. Louis, Economic Data-FRED. The i^{Long} is from the website of Board of Governors of the Federal Reserve. $i^{Germany}$ is from OECD.Stat Extracts. The NAIRU is constructed by the Congressional Budget Office and is downloaded from the Global Insight.

²³The GMM estimation is implemented based on the MATLAB package provided by Michael Cliff with documentation Cliff (2003).

(1.4) with annual data over the same time span for all the cases. As shown in Table 1.3, the qualitative result of the baseline estimation is not affected. Quantitatively, $\hat{\theta}_{GMM}$ with annual data approximately quadruples the magnitude of $\hat{\theta}_{GMM}$ of the baseline result. This pattern, which prevails in all the cases, may be explained by the time aggregation in the annual data. Compared with Claeys (2006), the result here with annual data indicates a much stronger response of primary surplus to lagged government debt.

1.5.3 The Bivariate VAR

Following Canzoneri, Cumby and Diba (2001), I estimate and identify a bivariate VAR to draw inferences on the U.S. fiscal policy behavior. The data include primary surplus/GDP and total government liabilities/GDP. Unlike the single-equation approaches, e.g. OLS and GMM, the bivariate VAR focuses on the impulse response of liabilities/GDP to a positive surplus/GDP shock.

The impulse responses of surplus/GDP and liabilities/GDP to a positive surplus/GDP shock are shown in Figure 1.6. The initial shock is assumed to be 1% of the steady-state surplus/GDP. The impulse response of each variable is scaled by its steady-state value. The baseline estimation result, which is based on the U.S. quarterly data over the sample 1947:2-2006:2, is shown in the top panel of Figure 1.6. The VAR features three lags and a constant. The variables in the VAR are ordered as surplus/GDP and liabilities/GDP. As in Canzoneri, Cumby and Diba (2001), identification of the structural VAR is based on Cholesky decomposition.

Clearly, qualitative result of the baseline estimation is consistent with the dynamics shown in the top panel of Figure 3 of Canzoneri, Cumby and Diba (2001). Particularly, a positive surplus/GDP shock induces a significant increase in surplus/GDP, which lasts for more than one year on 90% level. The liabilities/GDP significantly decreases on impact and keeps declining for a long time, i.e. about three years on 90% level. According to the argument of Canzoneri, Cumby and Diba (2001), the impulse responses of the baseline result favor a Ricardian interpretation of the U.S. data more than a non-Ricardian one.

To check robustness of the baseline result, I estimate and identify a VAR using the U.S. annual data over the same time span, i.e. 1947-2006. The VAR features five lags and a constant. The estimation result is shown in the bottom panel of Figure 1.6. Basically, the pattern of impulse responses of the baseline result is not affected. The only difference is the decrease of liabilities/GDP, which is significant for about ten years on both 68% and 90% levels. This is much longer than the baseline result but is closer to the result of Canzoneri, Cumby and Diba (2001), i.e. the top-right panel of Figure 3 of that paper.

1.6 Empirical Analysis: A Simultaneous Equations Approach

From the previous illustration, it is clear that the simultaneity bias problem is prevailing when a simple fiscal policy rule is estimated in isolation from the IEC and

the monetary policy rule. In this section, a fundamental solution to the problem is explored. Theoretically, the problem can be solved by estimating a fiscal policy rule jointly with the IEC and the monetary policy rule in the simultaneous equations framework. However, the implementation of the ideal solution is not straightforward because the IEC is a present-value relation that contains an infinite sum. As an alternative, I take the IEC indirectly to the estimation by using the methodology advocated in Chung and Leeper (2009). Specifically, I impose the cross-equation restrictions implied by the IEC on a VAR instead of estimating the IEC directly. The approach is applied to the postwar U.S. data. By adopting the identification scheme proposed by Blanchard and Perotti (2002) and extended by Perotti (2004), I recover the structural VAR that includes both the monetary and the fiscal policy rules and draw inferences on the postwar U.S. fiscal policy behavior. Since information of the IEC and the monetary policy behavior is embedded in the estimation, the approach is free of the simultaneity bias problem and the associated identification results are reliable.

In this section, I first review the necessary details of Chung and Leeper (2009)'s methodology for clarity. Then I describe the data set and the identification scheme of the VAR. Finally, I report the estimation results.

1.6.1 Government Budget Constraint and the IEC

First, I define the nominal market value of total government debt outstanding at time t , V_t , as

$$V_t \equiv \sum_{j=1}^{\infty} B_t(j) Q_t(j) \quad (1.29)$$

in which $B_t(j)$ is the nominal face value at time t of a zero coupon bond maturing at $t + j$ and $Q_t(j)$ is the market price of $B_t(j)$. I also define the nominal primary surplus, S_t , as

$$S_t = T_t - G_t - Z_t \quad (1.30)$$

in which T_t is the federal taxes, G_t is the government spending and Z_t is the net transfers. The government budget constraint is then given by

$$\sum_{j=1}^{\infty} [B_t(j) - B_{t-1}(j+1)] Q_t(j) = B_{t-1}(1) - S_t - (M_t - M_{t-1}) \quad (1.31)$$

in which M_t is the nominal money balance and $M_t - M_{t-1}$ is the nominal seigniorage revenues. It can be shown that the government budget constraint (1.31) is equivalent to

$$\frac{V_t}{P_t} = \frac{P_{t-1}}{P_t Q_{t-1}(1)} \frac{V_{t-1}}{P_{t-1}} - \frac{S_t}{P_t} - \frac{(M_t - M_{t-1})}{P_t} + \nu_t \quad (1.32)$$

in which P_t is the aggregate price level and

$$\nu_t \equiv \frac{1}{P_t} \sum_{j=1}^{\infty} \left[Q_t(j) - \frac{Q_{t-1}(j+1)}{Q_{t-1}(1)} \right] B_{t-1}(j+1) \quad (1.33)$$

Now I define the total nominal government liabilities as $W_t = V_t + M_t$. In terms of W_t , government budget constraint (1.32) can be written as

$$\frac{W_t}{P_t} = \frac{P_{t-1}}{P_t Q_{t-1}(1)} \frac{W_{t-1}}{P_{t-1}} - \frac{S_t}{P_t} - \left[\frac{P_{t-1}}{P_t Q_{t-1}(1)} - \frac{P_{t-1}}{P_t} \right] \frac{M_{t-1}}{P_{t-1}} + \nu_t \quad (1.34)$$

Let the one-period nominal interest rate at time $t - 1$ be $R_{t-1} = Q_{t-1}^{-1}(1)$. In real terms, government budget constraint (1.34) becomes

$$w_t = \frac{R_{t-1}}{\pi_t} w_{t-1} - s_t - \frac{1}{\pi_t} (R_{t-1} - 1) m_{t-1} + \nu_t \quad (1.35)$$

in which $w_t \equiv W_t/P_t$, $s_t \equiv S_t/P_t \equiv T_t/P_t - G_t/P_t - Z_t/P_t \equiv t_t - g_t - z_t \equiv \tau_t - g_t$, $m_t \equiv M_t/P_t$ and the inflation rate $\pi_t \equiv P_t/P_{t-1}$. Note, I define the taxes net of transfers (or net taxes) as $\tau_t \equiv t_t - z_t$.

To derive the IEC, I assume there is an equilibrium model generating the data and the Euler equation for the nominal zero coupon bond is given by

$$Q_t(j) = \delta^j E_t \frac{\lambda_{t+j}}{\lambda_t} \frac{P_t}{P_{t+j}} \quad (1.36)$$

where $\delta \in (0, 1)$ is the subjective discount rate and λ_t is the Lagrangian multiplier of the associated household problem. Applying (1.36) in (1.33), I get

$$\nu_t \equiv \frac{1}{\lambda_t} \underbrace{\sum_{j=1}^{\infty} \delta^j \left(E_t \frac{\lambda_{t+j}}{P_{t+j}} - \frac{\lambda_t}{P_t} \frac{E_{t-1} \frac{\lambda_{t+j}}{P_{t+j}}}{E_{t-1} \frac{\lambda_t}{P_t}} \right)}_{\eta_t} B_{t-1}(j+1) \quad (1.37)$$

or $\eta_t \equiv \lambda_t \nu_t$. It can be easily shown that, for $l \geq 1$,

$$E_{t+l-1}\eta_{t+l} = E_{t+l-1} \sum_{j=1}^{\infty} \delta^j \left(E_{t+l} \frac{\lambda_{t+l+j}}{P_{t+l+j}} - \frac{\lambda_{t+l}}{P_{t+l}} \frac{E_{t+l-1} \frac{\lambda_{t+l+j}}{P_{t+l+j}}}{E_{t+l-1} \frac{\lambda_{t+l}}{P_{t+l}}} \right) B_{t+l-1}(j+1) = 0$$

Hence, by the law of iterated expectations,

$$E_t \eta_{t+l} = 0, \text{ for } l \geq 1 \quad (1.38)$$

In equilibrium, the transversality condition for the household problem always holds as

$$\lim_{s \rightarrow \infty} E_t \lambda_{t+s} \frac{W_{t+s}}{P_{t+s}} = 0 \quad (1.39)$$

By iterating (1.34) forward, imposing (1.38) and (1.39) and taking expectation conditional on the information set at time t , I finally get the IEC, which, in real terms, is

$$w_t = E_t \sum_{j=1}^{\infty} \left(\prod_{k=1}^j \pi_{t+k} R_{t+k-1}^{-1} \right) [s_{t+j} + \pi_{t+j}^{-1} (R_{t+j-1} - 1) m_{t+j-1}] \quad (1.40)$$

For the purpose of linearizing the government budget constraint and the IEC around a stationary steady state, I scale all the real quantity variables in (1.35) and (1.40) by the real output $y_t \equiv Y_t/P_t$. Let the real output growth rate be $\mu_t \equiv \frac{Y_t/P_t}{Y_{t-1}/P_{t-1}}$. The scaled versions of (1.35) and (1.40) are then given by

$$w_t = \frac{R_{t-1}}{\pi_t \mu_t} w_{t-1} - s_t - \frac{1}{\pi_t \mu_t} (R_{t-1} - 1) m_{t-1} + \frac{P_t \nu_t}{Y_t} \quad (1.41)$$

and

$$w_t = E_t \sum_{j=1}^{\infty} \left(\prod_{k=1}^j \pi_{t+k} R_{t+k-1}^{-1} \mu_{t+k} \right) \left[s_{t+j} + (R_{t+j-1} - 1) \frac{m_{t+j-1}}{\pi_{t+j} \mu_{t+j}} \right] \quad (1.42)$$

in which w_t , s_t and m_t are ratios of the corresponding real quantity variables to the real output.

Next, I log-linearize the government budget constraint and the IEC around the stationary steady state and express all the variables in logarithm. In the stationary steady state, equilibrium condition (1.36) implies $\beta \equiv \pi\mu/R$, where β is the inverse of the steady state growth-adjusted real interest rate. The linearized version of (1.41) is given by

$$\begin{aligned} & \frac{\tau}{w} \tilde{\tau}_t - \frac{g}{w} \tilde{g}_t + \left[\frac{1}{\beta} - \frac{(R-1)m}{\pi\mu w} \right] \tilde{\mu}_t + \left[\frac{1}{\beta} - \frac{(R-1)m}{\pi\mu w} \right] \tilde{\pi}_t + \tilde{w}_t \\ = & k_1 + \frac{1}{\beta} \left(1 - \frac{m}{w} \right) \tilde{R}_{t-1} - \frac{(R-1)m}{\pi\mu w} \tilde{m}_{t-1} + \frac{1}{\beta} \tilde{w}_{t-1} + d \left(\frac{P_t \nu_t}{Y_t} \right) \end{aligned} \quad (1.43)$$

in which

$$\begin{aligned} k_1 \equiv & \frac{\tau}{w} \ln(\tau) - \frac{g}{w} \ln(g) + \left[\frac{1}{\beta} - \frac{(R-1)m}{\pi\mu w} \right] \ln(\mu) + \left[\frac{1}{\beta} - \frac{(R-1)m}{\pi\mu w} \right] \ln(\pi) \\ & - \frac{1}{\beta} \left(1 - \frac{m}{w} \right) \ln(R) + \frac{(R-1)m}{\pi\mu w} \ln(m) - \left(\frac{1}{\beta} - 1 \right) \ln(w) \end{aligned} \quad (1.44)$$

In (1.43), since the term $P_t \nu_t / Y_t$ can take negative values, I just take its deviation from the corresponding linearization point instead of log-linearizing it. Note, all the variables in (1.43) with a tilde are in logarithm and all the variables in (1.43) and

(1.44) without time subscript are steady state values. In particular, τ , g , m and w are the steady state ratios of the corresponding real quantity variables to the real output. Similarly, the linearized version of (1.42) is given by²⁴

$$\begin{aligned}\tilde{w}_t &= k_2 + E_t \sum_{j=1}^{\infty} \beta^j \left[\frac{\tau}{w} \tilde{\tau}_{t+j} - \frac{g}{w} \tilde{g}_{t+j} + \left(\kappa - \frac{(R-1)m}{\pi\mu w} \right) \tilde{\mu}_{t+j} \right] \\ &+ E_t \sum_{j=1}^{\infty} \beta^j \left[\left(\kappa - \frac{(R-1)m}{\pi\mu w} \right) \tilde{\pi}_{t+j} - \left(\kappa - \frac{Rm}{\pi\mu w} \right) \tilde{R}_{t+j-1} + \frac{(R-1)m}{\pi\mu w} \tilde{m}_{t+j-1} \right]\end{aligned}\quad (1.45)$$

in which $\kappa = \left[\frac{s}{w} + \frac{(R-1)m}{\pi\mu w} \right] \left(\frac{1}{1-\beta} \right)$ and

$$\begin{aligned}k_2 &\equiv \ln(w) - \frac{\beta}{1-\beta} \left[\kappa \ln(\beta) + \frac{\tau}{w} \ln(\tau) - \frac{g}{w} \ln(g) - \frac{(R-1)m}{\pi\mu w} \ln(\mu) \right] \\ &+ \frac{\beta}{1-\beta} \left[\frac{(R-1)m}{\pi\mu w} \ln(\pi) - \frac{Rm}{\pi\mu w} \ln(R) - \frac{(R-1)m}{\pi\mu w} \ln(m) \right]\end{aligned}\quad (1.46)$$

One novel feature of the approach proposed by Chung and Leeper (2009) is that maturity structure of government debt and term structure of interest rates are taken into account when the government budget constraint is derived. This has two important implications. First, the government budget constraint is more consistent with the real data than the conventional budget constraint that only includes one-period debt and short-term interest rate. Second, an error term, i.e. ν_t , is introduced to the government budget constraint and disappears in expectation. To derive the cross-equation restrictions from the IEC, all the variables in the IEC, which are also the variables in the government budget constraint, should be kept in the VAR.

²⁴See Appendix 1.C.1 for more details of the derivation of (1.45)-(1.46).

With the conventional government budget constraint in place, stochastic singularity problem may arise because the conventional budget constraint always holds as an identity. Fortunately, the problem is avoided by having ν_t in the government budget constraint.

1.6.2 Cross-Equation Restrictions on the VAR

Suppose the state of the economy is characterized by the n -dimensional factors f_t which evolve according to the VAR(p) process

$$f_t = B_0 + f_{t-1}B_1 + \cdots + f_{t-p}B_p + u_t \quad (1.47)$$

In the empirical model to be estimated, $n = 7$ and $f_t = [\tilde{\tau}_t, \tilde{g}_t, \tilde{\mu}_t, \tilde{\pi}_t, \tilde{R}_t, \tilde{m}_t, \tilde{w}_t]$ is a (1×7) vector, B_0 and u_t are $(1 \times n)$ vectors, B_i is an $(n \times n)$ matrix for $i = 1, \dots, p$. Also suppose all model variables \tilde{x}_t are related to the factors f_t through the mapping

$$\tilde{x}_t = f_t C_x \quad (1.48)$$

where, for example, the selection vector C_τ is simply defined as $C_\tau \equiv [1, 0, 0, 0, 0, 0, 0]'$.

Both f_t and \tilde{x}_t are measured in logarithm.

To derive the cross-equation restrictions from the IEC, it is convenient to fit (1.47) in companion form and express it as a VAR(1) process

$$\bar{f}_t = \bar{B}_0 + \bar{f}_{t-1}B + \bar{u}_t \quad (1.49)$$

in which $\bar{f}_t = [f_t, f_{t-1}, \dots, f_{t-p+1}]$, $\bar{B}_0 = [B_0, \mathbf{0}_{1 \times n(p-1)}]$ and $\bar{u}_t = [u_t, \mathbf{0}_{1 \times n(p-1)}]$ are all $(1 \times np)$ vectors and B is an $(np \times np)$ matrix which is defined as

$$B \equiv \begin{bmatrix} B_1 & \mathbf{I}_n & \mathbf{0}_n & \cdots & \mathbf{0}_n \\ B_2 & \mathbf{0}_n & \mathbf{I}_n & \cdots & \mathbf{0}_n \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ B_{p-1} & \mathbf{0}_n & \mathbf{0}_n & \cdots & \mathbf{I}_n \\ B_p & \mathbf{0}_n & \mathbf{0}_n & \cdots & \mathbf{0}_n \end{bmatrix}$$

in which \mathbf{I}_n is an n -dimensional identity matrix and $\mathbf{0}_n$ is an $(n \times n)$ zero square matrix. Accordingly, the mapping (1.48) becomes

$$\tilde{x}_t = \bar{f}_t \bar{C}_x \quad (1.50)$$

in which all \bar{C}_x are now $(np \times 1)$ vectors, e.g. $\bar{C}_\tau = [C'_\tau, \mathbf{0}'_{n(p-1) \times 1}]'$.

Appendix 1.C.2 shows that the IEC (1.45) implies the following cross-equation restrictions on the VAR coefficients \bar{B}_0 and B , which will be imposed in the estimation²⁵:

$$\begin{aligned} & \beta B \left\{ \frac{\tau}{w} \bar{C}_\tau - \frac{g}{w} \bar{C}_g + \left[\kappa - \frac{(R-1)m}{\pi \mu w} \right] \bar{C}_\mu + \left[\kappa - \frac{(R-1)m}{\pi \mu w} \right] \bar{C}_\pi + \bar{C}_w \right\} \\ &= \left(\kappa \beta - \frac{m}{w} \right) \bar{C}_R - \frac{(R-1)m}{Rw} \bar{C}_m + \bar{C}_w \end{aligned} \quad (1.51)$$

²⁵Note, only (1.51) is imposed in the current estimation. The imposition of (1.52), which implies restrictions on the deterministic growth components of the VAR, is left for the future research.

$$\begin{aligned}
& k_2 + \left(\frac{\beta}{1-\beta} \right) \bar{B}_0(I - \beta B)^{-1} \left[\frac{\tau}{w} \bar{C}_\tau - \frac{g}{w} \bar{C}_g + \left(\kappa - \frac{(R-1)m}{\pi\mu w} \right) \bar{C}_\mu \right] \\
& + \left(\frac{\beta}{1-\beta} \right) \bar{B}_0(I - \beta B)^{-1} \left[\left(\kappa - \frac{(R-1)m}{\pi\mu w} \right) \bar{C}_\pi + \left(\frac{m}{w} - \kappa\beta \right) \bar{C}_R \right] \\
& + \left(\frac{\beta}{1-\beta} \right) \bar{B}_0(I - \beta B)^{-1} \left[\frac{(R-1)m}{Rw} \bar{C}_m \right] = 0
\end{aligned} \tag{1.52}$$

1.6.3 Estimation Procedure

To estimate the unrestricted VAR without imposing (1.51), I apply the standard OLS regression. For the restricted VAR, I follow Chung and Leeper (2009) to apply a feasible generalized least squares (FGLS) procedure, which is briefly described as follows.

Suppose I want to estimate the reduced-form VAR(p) system (1.47)

$$f_t = B_0 + f_{t-1}B_1 + \cdots + f_{t-p}B_p + u_t \tag{1.53}$$

or more compactly

$$f_t = B_0 + f_- B_- + u_t \tag{1.54}$$

in which $f_- = [f_{t-1}, \dots, f_{t-p}]$ is a $(1 \times np)$ lagged data vector and $B_- = [B'_1, \dots, B'_p]'$ is an $(np \times n)$ matrix. Then I stack the $T + p$ observations in the form of (1.54) and get

$$F = \mathbf{1}B_0 + F_- B_- + u \tag{1.55}$$

in which F is a $(T \times n)$ data matrix, $\mathbf{1}$ is a $(T \times 1)$ vector consisting of all 1's, F_- is a $(T \times np)$ lagged data matrix and u is a $(T \times n)$ matrix. Next, I define

$X \equiv [\mathbf{1}, F_-]$, a $(T \times (np + 1))$ matrix, and $[B'_0, B'_-]'$, an $((np + 1) \times n)$ matrix. Let $b \equiv \text{vec}([B'_0, B'_-]')$, an $(n(np + 1) \times 1)$ vector. Applying vec operator, I can express (1.55) as

$$\text{vec}(F) = (\mathbf{I}_n \otimes X)b + \text{vec}(u) \quad (1.56)$$

The log-likelihood function corresponding to (1.56) is given by

$$\mathcal{L}(b, \Sigma) = -\frac{nT}{2} \ln(2\pi) + \frac{1}{2} \ln |(\Sigma \otimes I_T)^{-1}| - \frac{1}{2} [\text{vec}(u)'(\Sigma \otimes I_T)^{-1} \text{vec}(u)] \quad (1.57)$$

in which Σ is the cross-equation variance-covariance matrix.

The restriction (1.51) implied by the IEC can be simplified as

$$B_- C_1 = \bar{C}_0 \quad (1.58)$$

in which C_1 , an $(n \times 1)$ vector, and \bar{C}_0 , an $(np \times 1)$ vector, are given by

$$\begin{aligned} C_1 &\equiv \beta \left\{ \frac{\tau}{w} C_\tau - \frac{g}{w} C_g + \left[\kappa - \frac{(R-1)m}{\pi\mu w} \right] C_\mu + \left[\kappa - \frac{(R-1)m}{\pi\mu w} \right] C_\pi + C_w \right\} \\ \bar{C}_0 &\equiv \left(\kappa\beta - \frac{m}{w} \right) \bar{C}_R - \frac{(R-1)m}{Rw} \bar{C}_m + \bar{C}_w \end{aligned}$$

(1.58) can be rewritten as

$$[\mathbf{0}_{np \times 1}, \mathbf{I}_{np}][B'_0, B'_-]' C_1 = \bar{C}_0 \quad (1.59)$$

Applying vec operator on (1.59), I get

$$Vb = \text{vec}(\bar{C}_0) \quad (1.60)$$

in which $V \equiv (C_1' \otimes [\mathbf{0}_{np \times 1}, \mathbf{I}_{np}])$.

The goal of the estimation is to choose b to maximize (1.57) subject to the constraint (1.60). Appendix 1.C.3 shows that \tilde{b} , the consistent estimator of b of the restricted VAR, is given by

$$\tilde{b} = \hat{b} + S^{-1}V'(VS^{-1}V')^{-1}(\text{vec}(\bar{C}_0) - V\hat{b}) \quad (1.61)$$

in which $\hat{b} = \text{vec}((X'X)^{-1}(X'F))$ is the OLS estimator of b of the unrestricted VAR and $S \equiv (\Sigma^{-1} \otimes X'X)$. For implementation, I start with \hat{b} and use a FGLS procedure based on (1.61), iterated until convergence, to get consistent estimators of b and Σ of the restricted VAR. To obtain the distributions of the estimated parameters, I adopt the bias-adjusted bootstrap procedure proposed by Kilian (1998).

1.6.4 VAR Specification

Data

To estimate the VAR, I use the quarterly U.S. data from 1947:2 to 2006:2. The variables in logarithm are listed as follows: net taxes ($\tilde{\tau}_t$), government spending (\tilde{g}_t), real GDP growth rate ($\tilde{\mu}_t$), GDP deflator inflation rate ($\tilde{\pi}_t$), three-month Treasury

bill rate (\tilde{R}_t), monetary base (\tilde{m}_t) and total government liabilities (\tilde{w}_t). Note, all the real quantity variables in the VAR are scaled by the real GDP, which forms a cointegrated system. Appendix 1.C.4 provides more details of the data set construction, which follows Chung and Leeper (2009) and Traum (2007).

Identification Scheme

Assume there is a structural VAR,

$$f_t A = A_0 + f_{t-1} A_1 + \cdots + f_{t-p} A_p + \varepsilon_t J \quad (1.62)$$

which is a simultaneous equations system, underlying the reduced-form VAR (1.53).

Obviously, $B_i = A_i A^{-1}$ and $u_t = \varepsilon_t J A^{-1}$, in which u_t contains one-period ahead forecasting errors from the reduced-form VAR and ε_t contains structural shocks that are assumed to be orthogonal to each other.

To identify fiscal policy behavior, I need to recover ε_t from u_t , which amounts to identifying the contemporaneous matrix A . For this purpose, I adopt the identification scheme originally proposed by Blanchard and Perotti (2002) and extended by Perotti (2004). First, the block of fiscal variables is ordered before the other variables. Specifically, I assume

$$\begin{aligned} u_t^\tau &= \alpha_{\tau\mu} u_t^\mu + \alpha_{\tau\pi} u_t^\pi + \alpha_{\tau R} u_t^R + J_{\tau g} \varepsilon_t^g + \varepsilon_t^\tau \\ u_t^g &= \alpha_{g\mu} u_t^\mu + \alpha_{g\pi} u_t^\pi + \alpha_{gR} u_t^R + J_{g\tau} \varepsilon_t^\tau + \varepsilon_t^g \end{aligned} \quad (1.63)$$

in which all the α 's are elasticities that capture the automatic stabilizer embedded in the fiscal policy. The α 's in this paper are constructed based on the results of the previous literature²⁶, e.g. Perotti (2004) and Chung and Leeper (2009). Other than the automatic stabilizer, I assume fiscal policy does not respond to the state of the economy discretionarily within a quarter. According to this recursive ordering, the first two equations of the system represent the fiscal policy rules. Note, $J_{\tau g}$ and $J_{g\tau}$ in (1.63) are from the matrix J , which is a block diagonal matrix with the first block at the upper-left corner a two-dimensional matrix corresponding to the two fiscal shocks. For the estimation of this paper, I follow Chung and Leeper (2009) and choose the orthogonalization so that net taxes is ordered the first in the system²⁷, i.e. $J_{\tau g} = 0$.

In order to identify the whole matrix A , I assume the recursive ordering of the remaining variables as $[\tilde{\mu}_t, \tilde{\pi}_t, \tilde{R}_t, \tilde{m}_t, \tilde{w}_t]$, from left to right. For example, I assume

$$u_t^\mu = \alpha_{\mu\tau} u_t^\tau + \alpha_{\mu g} u_t^g + \varepsilon_t^\mu \quad (1.64)$$

Then I treat the identified fiscal shocks, $(\hat{\varepsilon}_t^\tau, \hat{\varepsilon}_t^g)$, as instruments of (u_t^τ, u_t^g) and run IV estimation on (1.64). Consistent estimates of $(\alpha_{\mu\tau}, \alpha_{\mu g})$ are obtained in this way. Equations like (1.64) for the other four variables can be written down recursively and the other elements in A can be identified accordingly.

²⁶Appendix 1.C.5 provides details for constructing the elasticities.

²⁷As a robustness check, I also try another orthogonalization, i.e. $J_{g\tau} = 0$. It turns out that the choice of orthogonalization does not affect the results much, which is consistent with the earlier finding in Blanchard and Perotti (2002) and Perotti (2004).

1.6.5 Estimation Results

Finally, I apply the simultaneous equations approach to identify the U.S. fiscal authority's stance on debt stabilization, which is free of the simultaneity bias problem a priori. Based on the quarterly U.S. data from 1947:2 to 2006:2, the baseline estimation features a VAR with two lags and a constant according to Bayesian information criterion (BIC). The steady-state values for the FGLS procedure are calibrated at sample means of the data set. To estimate distributions of the model parameters, the bias-adjusted bootstrap procedure proposed by Kilian (1998) is adopted.

To calculate the long-run response of surplus/GDP to lagged government debt/GDP, I start by considering a marginal increase in liabilities/GDP from the steady state, i.e. $\Delta w_t > 0$. Since $w_t = v_t + m_t$, I assume that part of Δw_t is brought about by Δv_t and the remaining part is due to Δm_t . I further assume that, on average, $\Delta v_t/\Delta w_t$ and $\Delta m_t/\Delta w_t$ are given by the steady-state ratios v/w and m/w , respectively. Note,

$$\frac{\Delta s}{\Delta w} = \frac{s}{w} LRE^{sw} \quad \text{and} \quad \frac{\Delta s}{\Delta m} = \frac{s}{m} LRE^{sm}$$

in which LRE^{sw} and LRE^{sm} are the long-run elasticities of surplus/GDP to liabilities/GDP and monetary base/GDP, respectively. Recalling that $s_t = \tau_t - g_t$, I

calculate LRE^{sw} and LRE^{sm} as follows

$$\begin{aligned}
LRE^{sw} &\equiv \sum_{j=1}^{\infty} \frac{d\tilde{s}_t}{d\tilde{w}_{t-j}} = \underbrace{\frac{\tau}{s} \sum_{j=1}^{\infty} \frac{d\tilde{\tau}_t}{d\tilde{w}_{t-j}}}_{LRE^{\tau w}} - \underbrace{\frac{g}{s} \sum_{j=1}^{\infty} \frac{d\tilde{g}_t}{d\tilde{w}_{t-j}}}_{LRE^{gw}} \\
LRE^{sm} &\equiv \sum_{j=1}^{\infty} \frac{d\tilde{s}_t}{d\tilde{m}_{t-j}} = \underbrace{\frac{\tau}{s} \sum_{j=1}^{\infty} \frac{d\tilde{\tau}_t}{d\tilde{m}_{t-j}}}_{LRE^{\tau m}} - \underbrace{\frac{g}{s} \sum_{j=1}^{\infty} \frac{d\tilde{g}_t}{d\tilde{m}_{t-j}}}_{LRE^{gm}}
\end{aligned}$$

in which $LRE^{\tau w}$ and $LRE^{\tau m}$ are the long-run elasticities of net taxes/GDP to liabilities/GDP and monetary base/GDP, respectively. Likewise, LRE^{gw} and LRE^{gm} are the long-run elasticities of spending/GDP to liabilities/GDP and monetary base/GDP, respectively. Note, $[LRE^{\tau w}, LRE^{\tau m}, LRE^{gw}, LRE^{gm}]$ are computed based on the first two equations of the identified VAR. Therefore, the long-run response of surplus/GDP to lagged government debt/GDP is given by

$$\frac{\Delta s}{\Delta v} = \frac{\Delta w * s/w * LRE^{sw} - \Delta w * m/w * s/m * LRE^{sm}}{\Delta w * v/w} = \frac{s}{v} LRE^{sw} - \frac{s}{v} LRE^{sm} \quad (1.65)$$

in which Δs is the long-run deviation of surplus/GDP from the steady state brought about by the marginal increase in lagged government debt/GDP, i.e. Δv .

The point estimate of $\Delta s/\Delta v$ from the baseline estimation is 0.0039, with the 90% confidence interval $[-0.3980, 0.5574]$. In terms of annual rate, $\Delta s/\Delta v$ is approximately 0.0155, which is less than half of the magnitude obtained from the OLS regression. In terms of quarterly rate, $\Delta s/\Delta v$ is about one-thirteenth of the magnitude obtained from the baseline GMM estimation. In a word, the simultaneous equations approach estimates a much weaker response of surplus/GDP to lagged

government debt/GDP in the long run. Besides, the interval estimate of $\Delta s/\Delta v$ is much wider than those reported by the single-equation approaches²⁸. This implies that the true value of $\Delta s/\Delta v$ is very uncertain. More importantly, the interval estimate of $\Delta s/\Delta v$ contains zero, which means that the long-run response is not significantly different from zero on 90% level.

From the baseline estimation, I also identify the monetary policy behavior by calculating the long-run response of nominal interest rate to inflation, i.e. $\Delta R/\Delta \pi$. It can be easily computed from the monetary policy rule, which is the fifth equation of the identified VAR. The point estimate of $\Delta R/\Delta \pi$ is 1.3143, with the 90% confidence interval [0.4011, 2.4725]. According to the point estimates of $\Delta R/\Delta \pi$ and $\Delta s/\Delta v$, both the monetary and the fiscal policy are inferred as being active, which is not consistent with our prior belief that the underlying dynamics of the economy is stationary.

Following Chung and Leeper (2009), I also investigate the forward-looking aspect of fiscal financing. In particular, I calculate the present-value components of government debt based on (1.45). The baseline result is shown in Table 1.4. The present-value decomposition shows what combination of adjustments in the expected paths of fiscal policy instruments, discount rates and seigniorage rationalizes the observed market value of government debt. Fiscal shocks are normalized to raise debt/GDP

²⁸With the simultaneous equations approach, the data-generating process is completely specified and is simulated using bootstrap method. This procedure accounts for the small-sample uncertainty in the parameters. While with the single-equation approaches, the advantage of the simultaneous equations approach mentioned earlier becomes infeasible and the confidence intervals are computed based on the asymptotic standard errors, which underestimate the small-sample uncertainty of the parameters. This may, to some extent, explain the huge difference in the width of the confidence intervals provided by different approaches.

from the steady state by 1%. From the baseline result, it is clear that future adjustments in fiscal policy depend on the nature of the fiscal shock. Specifically, the primary surplus component moves to support debt only when the debt innovation is financed by a spending hike. With a net taxes cut, the primary surplus component is offsetting the movement in debt, possibly because of the persistence in net taxes. Note, the movement of primary surplus is insignificant in both cases.

In summary, the long-run response of primary surplus to lagged government debt is insignificant if a fiscal policy rule is jointly estimated with the IEC and the monetary policy rule. This result holds both in terms of $\Delta s/\Delta v$ and from the forward-looking perspective of fiscal financing. Compared with the single-equation approaches such as OLS and GMM, the empirical implications on the U.S. fiscal policy behavior drawn by the simultaneous equations approach is qualitatively different.

As a contrast to the bivariate VAR, I compute the impulse responses of surplus/GDP and liabilities/GDP based on the simultaneous equations approach. Fiscal shocks are normalized to raise surplus/GDP from the steady state by 1%. The baseline result is shown in Figure 1.7, from which it is clear that the dynamics of the variables show different patterns depending on the nature of the fiscal shock. In the left panel of Figure 1.7, the 1% increase in surplus/GDP that is financed by a surprise net taxes hike is only significantly positive on impact. The dynamics of surplus/GDP is U-shaped, which is insignificant in the first quarter after the initial shock and then becomes significantly negative for 17 quarters on 68% level and 11 quarters on 90% level. For the liabilities/GDP, the median response on impact is

to increase by 0.01%, which is significant on both 68% and 90% levels. The dynamics is hump-shaped because of the U-shaped dynamics of surplus/GDP. Also, the impulse response of liabilities/GDP remains significantly positive for 10 years on 68% level and 35 quarters on 90% level. When the 1% increase in surplus/GDP is financed by a surprise spending cut, as shown in the right panel of Figure 1.7, the surplus/GDP increases a little 1 quarter later and then persistently declines. It remains significantly positive for 3 quarters on 68% level and only 1 quarter on 90% level. Following the spending shock, the median response of liabilities/GDP on impact is to decrease insignificantly by 0.0023%. It then remains insignificant for 10 years on both 68% and 90% levels.

Obviously, after imposing the cross-equation restrictions derived from the IEC on the unrestricted VAR that includes both the monetary and the fiscal policy rules, the patterns of the impulse responses observed in the bivariate VAR are overturned. With the simultaneous equations approach, the increase in surplus/GDP is very short-lived. The contemporaneous response of liabilities/GDP is either significantly positive or insignificant, which can hardly be explained by the Ricardian interpretation as argued in Canzoneri, Cumby and Diba (2001).

To check the robustness of the baseline results, some sensitivity analysis is carried out. First, according to Akaike information criterion (AIC), a VAR with three lags and a constant is estimated over the same sample as in the baseline estimation. Second, considering the critical role played by the elasticities in (1.63), a proper range of values for the elasticities is explored besides the particular set of values

calculated in Appendix 1.C.5. Third, the simultaneous equations approach is applied to different samples over which the OLS and GMM are estimated in the previous sections. It turns out that making changes in these aspects does not affect the qualitative results of the baseline estimation²⁹.

1.7 Conclusion

In order to identify the fiscal authority's stance on debt stabilization, the existing limited-information approaches, such as the OLS, GMM and bivariate VAR, only focus on estimating the fiscal policy rules, but fail to take the IEC of debt valuation and the monetary policy behavior into account. This paper shows that, in general, ignoring the IEC and the monetary policy behavior in the estimation results in a simultaneity bias problem and, therefore, the corresponding inferences may be misleading. The paper also shows that dealing with the specification of the fiscal policy rules (e.g., introducing policy inertia) or the fiscal shock processes is not a robust solution to the simultaneity bias problem. In order to solve the problem fundamentally, I propose a simultaneous equations approach that incorporates the IEC and the monetary policy rule in the estimation. Based on the post-World War II U.S. data, I find out that, on average, a non-Ricardian interpretation of the U.S. data is at least as plausible as a Ricardian one, which is completely different from the results estimated by the OLS, GMM and bivariate VAR methods.

In this paper, jointly identifying both monetary and fiscal policy behavior is highly

²⁹The estimation results of the sensitivity analysis is available from the author upon request.

emphasized, because it not only avoids the simultaneity bias problem but also provides a better understanding of the nature of the economy. However, with both kinds of policy behavior being inferred as active, the empirical results of this paper leaves this question as a puzzle. Several things deserve further investigation in the future work. First, only the point estimates are used to draw inferences on the joint policy behavior in the current work. Obviously, the estimation uncertainty needs to be taken into account as well. To accomplish this, Bayesian VAR may be considered as a way to go. Second, the prior belief that the underlying dynamics of the economy is stationary may not be right. So it is worthwhile to explore the econometric methodology for a non-stationary VAR in the future. Last but not the least, the monetary and the fiscal policy rules estimated in this paper are far more complicated than the simple ones in the current literature. So the current criteria to determine “active” or “passive”, which are originally proposed for the simple policy rules, may be inappropriate for the complicated policy rules in this paper. How to distinguish different policy behavior based on complicated policy rules is still unclear in the literature, which is also beyond the scope of this paper.

Clearly, the simultaneous equations approach only identifies the averaged fiscal policy behavior. According to the recent literature³⁰, the U.S. post-World War II fiscal regime was extremely variable. More precisely, the U.S. fiscal policy between the 1960s and the 1980s may be characterized as active. It then gradually switched to passive in the early 1990s and switched back to active in early 2001. Therefore, to better identify the real fiscal policy behavior, regime-switching fiscal policy rules

³⁰Examples include Favero and Monacelli (2003, 2005), Sala (2004), Davig and Leeper (2006) and Davig, Leeper and Chung (2007), among others.

should be used, which will be explored in the future.

Appendix 1.A: The Illustrative DSGE Model

Appendix 1.A.1: Model Solution

To solve the model, I derive the first-order necessary conditions of the household problem. Together with the aggregate resource constraint (1.8), I obtain the Fisher equation and the money-demand relation as follows:

$$\frac{1}{R_t} = \beta E_t \left[\frac{1}{\pi_{t+1}} \right] \quad (1.66)$$

$$m_t = \delta c \left[\frac{R_t}{R_t - 1} \right] \quad (1.67)$$

in which $m_t \equiv M_t/P_t$ is the real money balance.

Next, I log-linearize (1.66) and (1.9) as

$$\hat{R}_t = E_t \hat{\pi}_{t+1} \quad (1.68)$$

$$\hat{R}_t = \alpha \hat{\pi}_t + \hat{\theta}_t \quad (1.69)$$

in which I impose the steady state condition $R = \beta^{-1}$ and assume $\pi = 1$ for a zero-inflation steady state. Note, \hat{x}_t denotes log deviation of the generic variable x_t from its steady state value x , i.e. $\hat{x}_t \equiv \ln(x_t) - \ln(x)$. Combining (1.68) and (1.69),

I get

$$E_t \hat{\pi}_{t+1} = \alpha \hat{\pi}_t + \hat{\theta}_t \quad (1.70)$$

I then define the one-period-ahead endogenous forecasting error $\eta_{t+1} \equiv \hat{\pi}_{t+1} - E_t \hat{\pi}_{t+1}$ and express (1.70) as

$$\hat{\pi}_{t+1} = \alpha \hat{\pi}_t + \hat{\theta}_t + \eta_{t+1} \quad (1.71)$$

Next, I log-linearize (2.27), (1.67) and (1.10) as

$$m\hat{m}_t + b\hat{b}_t + \tau\hat{\tau}_t = m\hat{m}_{t-1} - m\hat{\pi}_t + Rb\hat{R}_{t-1} + Rb\hat{b}_{t-1} - Rb\hat{\pi}_t \quad (1.72)$$

$$m(R-1)\hat{m}_t = R(\delta c - m)\hat{R}_t \quad (1.73)$$

$$\hat{\tau}_t = \gamma\hat{b}_{t-1} + \hat{\psi}_t \quad (1.74)$$

in which I impose the relevant steady state conditions. Combining (1.69), (2.37)-(1.74) and rearranging terms, I get

$$\varphi_1 \hat{\pi}_t + \hat{b}_t + \varphi_2 \hat{\pi}_{t-1} - [\beta^{-1} - \gamma(\beta^{-1} - 1)] \hat{b}_{t-1} + \varphi_3 \hat{\theta}_t + (\beta^{-1} - 1) \hat{\psi}_t + \varphi_4 \hat{\theta}_{t-1} = 0 \quad (1.75)$$

in which

$$\varphi_1 \equiv \frac{m}{b}(\alpha\chi + 1) + \beta^{-1}$$

$$\varphi_2 \equiv -\alpha \left[\frac{m}{b}\chi + \beta^{-1} \right]$$

$$\varphi_3 \equiv \frac{m}{b}\chi$$

$$\varphi_4 \equiv - \left[\frac{m}{b}\chi + \beta^{-1} \right]$$

and $\chi \equiv 1/(1-R)$ is the negative interest elasticity of money demand. It is straightforward to log-linearize (1.11) and (1.12) as

$$\hat{\theta}_t = \rho_\theta \hat{\theta}_{t-1} + \varepsilon_t^\theta \quad (1.76)$$

$$\hat{\psi}_t = \rho_\psi \hat{\psi}_{t-1} + \varepsilon_t^\psi \quad (1.77)$$

After substituting (1.76) and (1.77) into (1.75) and expressing terms one period forwards, I get

$$\begin{aligned} \varphi_1 \hat{\pi}_{t+1} + \hat{b}_{t+1} = & -\varphi_2 \hat{\pi}_t + [\beta^{-1} - \gamma(\beta^{-1} - 1)] \hat{b}_t - (\varphi_3 \rho_\theta + \varphi_4) \hat{\theta}_t - (\beta^{-1} - 1) \rho_\psi \hat{\psi}_t \\ & - \varphi_3 \varepsilon_{t+1}^\theta - (\beta^{-1} - 1) \varepsilon_{t+1}^\psi \end{aligned} \quad (1.78)$$

So far, (1.71), (1.76)-(1.78) form a self-contained system governing the dynamics of $\hat{\pi}_t$, \hat{b}_t , $\hat{\theta}_t$ and $\hat{\psi}_t$. The linearized system is organized in the following compact form

$$\Gamma_0 Y_{t+1} = \Gamma_1 Y_t + \Pi \eta_{t+1} + \Psi \varepsilon_{t+1} \quad (1.79)$$

in which $Y_{t+1} = [\hat{\pi}_{t+1}, \hat{b}_{t+1}, \hat{\theta}_{t+1}, \hat{\psi}_{t+1}]'$, $\varepsilon_{t+1} = [\varepsilon_{t+1}^\theta, \varepsilon_{t+1}^\psi]'$ and

$$\Gamma_0 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ \varphi_1 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}, \quad \Pi = \begin{bmatrix} 1 \\ 0 \\ 0 \\ 0 \end{bmatrix}, \quad \Psi = \begin{bmatrix} 0 & 0 \\ -\varphi_3 & -(\beta^{-1} - 1) \\ 1 & 0 \\ 0 & 1 \end{bmatrix}$$

$$\Gamma_1 = \begin{bmatrix} \alpha & 0 & 1 & 0 \\ -\varphi_2 & \beta^{-1} - \gamma(\beta^{-1} - 1) & -(\varphi_3\rho_\theta + \varphi_4) & -(\beta^{-1} - 1)\rho_\psi \\ 0 & 0 & \rho_\theta & 0 \\ 0 & 0 & 0 & \rho_\psi \end{bmatrix}$$

Since Γ_0 is invertible, (1.79) can be written as

$$Y_{t+1} = \Gamma_1^* Y_t + \Pi^* \eta_{t+1} + \Psi^* \varepsilon_{t+1}$$

in which $\Gamma_1^* = \Gamma_0^{-1} \Gamma_1$, $\Pi^* = \Gamma_0^{-1} \Pi$ and $\Psi^* = \Gamma_0^{-1} \Psi$. Applying a Jordan decomposition on Γ_1^* , the equation above becomes:

$$Y_{t+1} = P \Lambda P^{-1} Y_t + \Pi^* \eta_{t+1} + \Psi^* \varepsilon_{t+1} \quad (1.80)$$

in which Λ is a diagonal matrix with the eigenvalues of Γ_1^* on the main diagonal, i.e.

$$\Lambda = \begin{bmatrix} \alpha & 0 & 0 & 0 \\ 0 & \beta^{-1} - \gamma(\beta^{-1} - 1) & 0 & 0 \\ 0 & 0 & \rho_\theta & 0 \\ 0 & 0 & 0 & \rho_\psi \end{bmatrix}$$

and P is a matrix composed of the right eigenvectors of Γ_1^* , in which each column is the eigenvector associated with the corresponding eigenvalue in Λ . It is well known that the determinacy of bounded equilibrium of the linearized model hinges on the

eigenvalues of Γ_1^* , which are $[\alpha, \beta^{-1} - \gamma(\beta^{-1} - 1), \rho_\theta, \rho_\psi]$. In fact, the determinacy solely hinges on (α, γ) , the two parameters in the monetary and the fiscal policy rules. When Γ_1^* does not have repeated eigenvalues, P has full column rank and (1.80) can be written as

$$P^{-1}Y_{t+1} = \Lambda P^{-1}Y_t + P^{-1}\Pi^*\eta_{t+1} + P^{-1}\Psi^*\varepsilon_{t+1} \quad (1.81)$$

in which P^{-1} can be shown as

$$P^{-1} = \begin{bmatrix} \frac{\alpha\varphi_1+\varphi_2}{\beta^{-1}-\gamma(\beta^{-1}-1)-\alpha} & 0 & \frac{\alpha\varphi_1+\varphi_2}{[\beta^{-1}-\gamma(\beta^{-1}-1)-\alpha](\alpha-\rho_\theta)} & 0 \\ -\frac{\alpha\varphi_1+\varphi_2}{\beta^{-1}-\gamma(\beta^{-1}-1)-\alpha} & 1 & P_{2,3}^{-1} & -\frac{(\beta^{-1}-1)\rho_\psi}{\beta^{-1}-\gamma(\beta^{-1}-1)-\rho_\psi} \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$$

in which $P_{2,3}^{-1} = -\frac{\varphi_1[\beta^{-1}-\gamma(\beta^{-1}-1)]+\varphi_2}{[\beta^{-1}-\gamma(\beta^{-1}-1)-\alpha][\beta^{-1}-\gamma(\beta^{-1}-1)-\rho_\theta]} - \frac{\varphi_3\rho_\theta+\varphi_4}{\beta^{-1}-\gamma(\beta^{-1}-1)-\rho_\theta}$. When Γ_1^* has repeated eigenvalues, as in the simple case I and II of section 3, P does not have full column rank and generalized eigenvectors are obtained.

If the bounded equilibrium is determinate, the following conditions suppress the unstable root of the dynamic system and hold for all t :

$$P^{i\cdot}Y_t = 0 \quad (1.82)$$

$$P^{i\cdot}\Pi^*\eta_{t+1} + P^{i\cdot}\Psi^*\varepsilon_{t+1} = 0 \quad (1.83)$$

in which $P^{i\cdot}$ is the i th row of P^{-1} and i is the index denoting the unstable eigenvalue

in Λ . In this paper, $i = 1$ is corresponding to AM/PF and 2 is corresponding to PM/AF. In both cases, η_{t+1} is uniquely determined by (1.83). If the bounded equilibrium is indeterminate, there is no unique mapping between η_{t+1} and ε_{t+1} .

Appendix 1.A.2: Simultaneity Bias in the Simple Cases

- **Case III (PM/AF):** $\gamma = \rho_\theta = 0$, $0 < \alpha < 1$, $0 < \rho_\psi < 1$

In this case, condition (1.82) is given by

$$a_{21}\hat{\pi}_t + \hat{b}_t + a_{23}\hat{\theta}_t + a_{24}\hat{\psi}_t = 0 \quad (1.84)$$

in which a_{ij} is the ij th entry of P^{-1} . Since $a_{21} \neq 0$, I substitute (1.84) into (1.75) and get

$$\begin{aligned} & \underbrace{\left(\frac{\varphi_1}{a_{21}} - 1\right)}_{A_1} \hat{b}_t + \underbrace{\left[\frac{\varphi_2}{a_{21}} + \beta^{-1} - \gamma(\beta^{-1} - 1)\right]}_{A_2} \hat{b}_{t-1} = \underbrace{\left(\varphi_3 - \frac{\varphi_1 a_{23}}{a_{21}}\right)}_{A_3} \hat{\theta}_t \\ & + \underbrace{\left(\beta^{-1} - 1 - \frac{\varphi_1 a_{24}}{a_{21}}\right)}_{A_4} \hat{\psi}_t + \underbrace{\left(\varphi_4 - \frac{\varphi_2 a_{23}}{a_{21}}\right)}_{A_5} \hat{\theta}_{t-1} + \underbrace{\left(-\frac{\varphi_2 a_{24}}{a_{21}}\right)}_{A_6} \hat{\psi}_{t-1} \end{aligned}$$

Since $A_1 \neq 0$, I have

$$\begin{aligned}
\hat{b}_t &= \frac{A_3 + A_5 L}{(A_1 + A_2 L)(1 - \rho_\theta L)} \varepsilon_t^\theta + \frac{A_4 + A_6 L}{(A_1 + A_2 L)(1 - \rho_\psi L)} \varepsilon_t^\psi \\
&= \frac{\frac{A_3}{A_1} + \frac{A_5}{A_1} L}{(1 + \frac{A_2}{A_1} L)(1 - \rho_\theta L)} \varepsilon_t^\theta + \frac{\frac{A_4}{A_1} + \frac{A_6}{A_1} L}{(1 + \frac{A_2}{A_1} L)(1 - \rho_\psi L)} \varepsilon_t^\psi \\
&= \underbrace{\frac{C_3 + C_5 L}{(1 + C_2 L)(1 - \rho_\theta L)}}_{B_1} \varepsilon_t^\theta + \underbrace{\frac{C_4 + C_6 L}{(1 + C_2 L)(1 - \rho_\psi L)}}_{B_2} \varepsilon_t^\psi
\end{aligned}$$

in which $C_i = A_i/A_1$, for $i = 2, 3, 4, 5, 6$.

It can be shown that

$$B_1 = \frac{m_1}{1 + C_2 L} - \frac{n_1}{1 - \rho_\theta L} \quad \text{and} \quad B_2 = \frac{m_2}{1 + C_2 L} - \frac{n_2}{1 - \rho_\psi L}$$

in which

$$m_1 = \frac{C_2 C_3 - C_5}{\rho_\theta + C_2}, \quad m_2 = \frac{C_2 C_4 - C_6}{\rho_\psi + C_2}, \quad n_1 = -\frac{C_3 \rho_\theta + C_5}{\rho_\theta + C_2}, \quad n_2 = -\frac{C_4 \rho_\psi + C_6}{\rho_\psi + C_2}$$

So I have

$$\begin{aligned}
\hat{b}_{t-1} &= B_1 \varepsilon_{t-1}^\theta + B_2 \varepsilon_{t-1}^\psi \\
&= (m_1 - n_1) \varepsilon_{t-1}^\theta - (m_1 C_2 + n_1 \rho_\theta) \varepsilon_{t-2}^\theta + (m_1 C_2^2 - n_1 \rho_\theta^2) \varepsilon_{t-3}^\theta - \cdots \\
&\quad + (m_2 - n_2) \varepsilon_{t-1}^\psi - (m_2 C_2 + n_2 \rho_\psi) \varepsilon_{t-2}^\psi + (m_2 C_2^2 - n_2 \rho_\psi^2) \varepsilon_{t-3}^\psi - \cdots
\end{aligned}$$

Since

$$\hat{\psi}_t = \frac{\varepsilon_t^\psi}{1 - \rho_\psi L} = \varepsilon_t^\psi + \rho_\psi \varepsilon_{t-1}^\psi + \rho_\psi^2 \varepsilon_{t-2}^\psi + \dots$$

I can show that

$$\begin{aligned} & \text{cov}(\hat{b}_{t-1}, \hat{\psi}_t) \\ &= (m_2 - n_2)\rho_\psi \sigma_\psi^2 - (m_2 C_2 + n_2 \rho_\psi)\rho_\psi^2 \sigma_\psi^2 + (m_2 C_2^2 - n_2 \rho_\psi^2)\rho_\psi^3 \sigma_\psi^2 - \dots \\ &= [(m_2 \rho_\psi - m_2 C_2 \rho_\psi^2 + m_2 C_2^2 \rho_\psi^3 - \dots) - (n_2 \rho_\psi + n_2 \rho_\psi^3 + n_2 \rho_\psi^5 + \dots)]\sigma_\psi^2 \\ &= \left(\frac{m_2 \rho_\psi}{1 + C_2 \rho_\psi} - \frac{n_2 \rho_\psi}{1 - \rho_\psi^2} \right) \sigma_\psi^2 \end{aligned} \tag{1.85}$$

in which the last equality holds because when monetary policy is passive, $C_2 = -\alpha\beta > -1$ and $|C_2 \rho_\psi| < 1$.

Finally, I can show that

$$\begin{aligned} & \text{var}(\hat{b}_{t-1}) \\ &= \text{cov}(\hat{b}_{t-1}, \hat{b}_{t-1}) \\ &= [(m_1 - n_1)^2 + (m_1 C_2 + n_1 \rho_\theta)^2 + (m_1 C_2^2 - n_1 \rho_\theta^2)^2 + \dots]\sigma_\theta^2 \\ & \quad + [(m_2 - n_2)^2 + (m_2 C_2 + n_2 \rho_\psi)^2 + (m_2 C_2^2 - n_2 \rho_\psi^2)^2 + \dots]\sigma_\psi^2 \\ &= \frac{m_1^2 \sigma_\theta^2 + m_2^2 \sigma_\psi^2}{1 - C_2^2} + \frac{n_1^2 \sigma_\theta^2}{1 - \rho_\theta^2} + \frac{n_2^2 \sigma_\psi^2}{1 - \rho_\psi^2} - \frac{2m_1 n_1 \sigma_\theta^2}{1 + C_2 \rho_\theta} - \frac{2m_2 n_2 \sigma_\psi^2}{1 + C_2 \rho_\psi} \end{aligned} \tag{1.86}$$

- **Case IV (AM/PF):** $\alpha > 1$, $\gamma > 1$, $0 < \rho_\psi < 1$

In this case, condition (1.82) and (1.83) simplify as

$$\hat{\pi}_t = -\frac{1}{\alpha - \rho_\theta} \hat{\theta}_t \quad (1.87)$$

$$\eta_t = -\frac{1}{\alpha - \rho_\theta} \varepsilon_t^\theta \quad (1.88)$$

from which it is clear that fiscal shock has no effect on inflation. This is the standard result of Ricardian equivalence, which is mentioned in the paper.

Since the DGP is characterized by AM/PF, the second row of (1.81), as defined in Appendix 1.A.1, is a stable first-order difference equation, from which I can solve for \hat{b}_t as a function of ε_t^θ and ε_t^ψ :

$$\begin{aligned} \hat{b}_t = & \frac{D_3 + D_5 L}{(D_1 + D_2 L)(1 - \rho_\theta L)} \varepsilon_t^\theta + \frac{D_4 + D_6 L}{(D_1 + D_2 L)(1 - \rho_\psi L)} \varepsilon_t^\psi \\ & + \frac{D_7}{D_1 + D_2 L} \varepsilon_t^\theta + \frac{D_8}{D_1 + D_2 L} \varepsilon_t^\psi \end{aligned}$$

in which

$$D_1 \equiv 1$$

$$D_2 \equiv -\lambda$$

$$D_3 \equiv -(a_{21}g + a_{23})$$

$$D_4 \equiv -a_{24}$$

$$D_5 \equiv \lambda(a_{21}g + a_{23})$$

$$D_6 \equiv \lambda a_{24}$$

$$D_7 \equiv \left(-\frac{\varphi_1\alpha + \varphi_2}{\lambda - \alpha} - \varphi_1 \right) g - \varphi_3 + a_{23}$$

$$D_8 \equiv -(\beta^{-1} - 1) - \frac{(\beta^{-1} - 1)\rho_\psi}{\lambda - \rho_\psi}$$

$$g \equiv -1/(\alpha - \rho_\theta)$$

$$\lambda \equiv \beta^{-1} - \gamma(\beta^{-1} - 1)$$

It can be shown that

$$\begin{aligned} \frac{D_3 + D_5 L}{(D_1 + D_2 L)(1 - \rho_\theta L)} &= \frac{m_1}{1 - \lambda L} - \frac{n_1}{1 - \rho_\theta L} \\ &= m_1 + \lambda m_1 L + \lambda^2 m_1 L^2 + \dots \\ &\quad - n_1 - \rho_\theta n_1 L - \rho_\theta^2 n_1 L^2 - \dots \\ \frac{D_4 + D_6 L}{(D_1 + D_2 L)(1 - \rho_\psi L)} &= \frac{m_2}{1 - \lambda L} - \frac{n_2}{1 - \rho_\psi L} \\ &= m_2 + \lambda m_2 L + \lambda^2 m_2 L^2 + \dots \\ &\quad - n_2 - \rho_\psi n_2 L - \rho_\psi^2 n_2 L^2 - \dots \end{aligned}$$

in which

$$m_1 = \frac{D_5 + \lambda D_3}{\lambda - \rho_\theta}, \quad m_2 = \frac{D_6 + \lambda D_4}{\lambda - \rho_\psi}, \quad n_1 = \frac{D_5 + \rho_\theta D_3}{\lambda - \rho_\theta}, \quad n_2 = \frac{D_6 + \rho_\psi D_4}{\lambda - \rho_\psi}$$

Besides,

$$\begin{aligned} \frac{D_7}{D_1 + D_2 L} &= D_7 + \lambda D_7 L + \lambda^2 D_7 L^2 + \dots \\ \frac{D_8}{D_1 + D_2 L} &= D_8 + \lambda D_8 L + \lambda^2 D_8 L^2 + \dots \end{aligned}$$

With some algebraic derivation, I can show that

$$\begin{aligned} \text{cov}(\hat{b}_{t-1}, \hat{\psi}_t) &= -\frac{\rho_\psi \sigma_\psi^2}{[1 - \lambda \rho_\psi](1 - \rho_\psi^2)} < 0 \\ \text{var}(\hat{b}_{t-1}) &= \left[\frac{(m_1 + D_7)^2}{1 - \lambda^2} + \frac{n_1^2}{1 - \rho_\theta^2} - \frac{2n_1(m_1 + D_7)}{1 - \lambda \rho_\theta} \right] \sigma_\theta^2 \\ &\quad + \left[\frac{(m_2 + D_8)^2}{1 - \lambda^2} + \frac{n_2^2}{1 - \rho_\psi^2} - \frac{2n_2(m_2 + D_8)}{1 - \lambda \rho_\psi} \right] \sigma_\psi^2 \end{aligned}$$

Appendix 1.B: The Model with Policy Inertia

To solve the model, I first combine (1.21), (1.69), (2.37) and (1.73). After rearranging terms, I get

$$\begin{aligned} & \varphi_1 \hat{\pi}_t + \hat{b}_t + \varphi_2 \hat{\pi}_{t-1} - [\beta^{-1} - \omega\gamma(\beta^{-1} - 1)] \hat{b}_{t-1} + (1 - \omega)(\beta^{-1} - 1) \hat{\tau}_{t-1} \\ & + \varphi_3 \hat{\theta}_t + (\beta^{-1} - 1) \hat{\psi}_t + \varphi_4 \hat{\theta}_{t-1} = 0 \end{aligned} \quad (1.89)$$

in which all the φ 's are defined as in Appendix 1.A.1. Next, I fit (1.21), (1.22), (1.71), (1.76) and (1.89) into a dynamic system in terms of $\hat{\pi}_t$, \hat{b}_t , $\hat{\tau}_t$, $\hat{\theta}_t$, $\hat{\psi}_t$ and ε_t^ψ .

The linearized system is organized as

$$\Gamma_0 Y_{t+1} = \Gamma_1 Y_t + \Pi \eta_{t+1} + \Psi \varepsilon_{t+1} \quad (1.90)$$

in which $Y_{t+1} = [\hat{\pi}_{t+1}, \hat{b}_{t+1}, \hat{\tau}_{t+1}, \hat{\theta}_{t+1}, \hat{\psi}_{t+1}, \varepsilon_{t+1}^\psi]'$, $\varepsilon_{t+1} = [\varepsilon_{t+1}^\theta, \varepsilon_{t+1}^\psi]'$ and

$$\Gamma_0 = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ \varphi_1 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}, \quad \Pi = \begin{bmatrix} 1 \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix}, \quad \Psi = \begin{bmatrix} 0 & 0 \\ -\varphi_3 & -(\beta^{-1} - 1)(1 - \phi) \\ 0 & 1 - \phi \\ 1 & 0 \\ 0 & 1 - \phi \\ 0 & 1 \end{bmatrix}$$

$$\Gamma_1 = \begin{bmatrix} \alpha & 0 & 0 & 1 & 0 & 0 \\ -\varphi_2 & \beta^{-1} - \omega\gamma(\beta^{-1} - 1) & -(1 - \omega)(\beta^{-1} - 1) & -(\varphi_3\rho_\theta + \varphi_4) & -(\beta^{-1} - 1)\rho_\psi & -(\beta^{-1} - 1)\phi \\ 0 & \omega\gamma & 1 - \omega & 0 & \rho_\psi & \phi \\ 0 & 0 & 0 & \rho_\theta & 0 & 0 \\ 0 & 0 & 0 & 0 & \rho_\psi & \phi \\ 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

To solve the model, I apply a Jordan decomposition as in Appendix 1.A.1. The matrix Λ is

$$\Lambda = \begin{bmatrix} \lambda_1 & 0 & 0 & 0 & 0 & 0 \\ 0 & \lambda_2 & 0 & 0 & 0 & 0 \\ 0 & 0 & \alpha & 0 & 0 & 0 \\ 0 & 0 & 0 & \rho_\theta & 0 & 0 \\ 0 & 0 & 0 & 0 & \rho_\psi & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

in which λ_1 and λ_2 are the two roots of the quadratic equation

$$\beta\lambda^2 + (\omega\beta - \beta - 1 + \omega\gamma - \omega\gamma\beta)\lambda + (1 - \omega) = 0 \quad (1.91)$$

As in the benchmark DSGE model, determinacy of the bounded equilibrium solely hinges on the policy parameters (α, γ) . The determinacy regions as in Figure 1.2 are still valid here. The remaining steps of the solution procedure follow Appendix 1.A.1.

Appendix 1.C: Simultaneous Equations Approach

Appendix 1.C.1: Linearizing the IEC

First, I rewrite (1.42) as

$$\begin{aligned}
 w_t &= E_t \sum_{j=1}^{\infty} \left(\prod_{k=1}^j \pi_{t+k} R_{t+k-1}^{-1} \mu_{t+k} \right) \left[s_{t+j} + (R_{t+j-1} - 1) \frac{m_{t+j-1}}{\pi_{t+j} \mu_{t+j}} \right] \\
 &= E_t \sum_{j=1}^{\infty} \left(\prod_{k=1}^j \pi_{t+k} R_{t+k-1}^{-1} \mu_{t+k} \right) s_{t+j} \\
 &\quad + E_t \sum_{j=1}^{\infty} \left(\prod_{k=1}^j \pi_{t+k} R_{t+k-1}^{-1} \mu_{t+k} \right) \frac{R_{t+j-1} m_{t+j-1}}{\pi_{t+j} \mu_{t+j}} \\
 &\quad - E_t \sum_{j=1}^{\infty} \left(\prod_{k=1}^j \pi_{t+k} R_{t+k-1}^{-1} \mu_{t+k} \right) \frac{m_{t+j-1}}{\pi_{t+j} \mu_{t+j}}
 \end{aligned} \tag{1.92}$$

Let $\Delta_t \equiv \prod_{k=1}^j \pi_{t+k} R_{t+k-1}^{-1} \mu_{t+k}$. Log linearizing (1.92) around the stationary steady state and rearranging terms, I get

$$\begin{aligned}
 \hat{w}_t &= E_t \sum_{j=1}^{\infty} \beta^j \left[\frac{s}{w} + \frac{(R-1)m}{\pi \mu w} \right] \hat{\Delta}_t + E_t \sum_{j=1}^{\infty} \beta^j \left[\frac{\tau}{w} \hat{\tau}_{t+j} - \frac{g}{w} \hat{g}_{t+j} \right] \\
 &\quad - E_t \sum_{j=1}^{\infty} \beta^j \left[\frac{(R-1)m}{\pi \mu w} \hat{\mu}_{t+j} + \frac{(R-1)m}{\pi \mu w} \hat{\pi}_{t+j} \right] \\
 &\quad + E_t \sum_{j=1}^{\infty} \beta^j \left[\frac{Rm}{\pi \mu w} \hat{R}_{t+j-1} + \frac{(R-1)m}{\pi \mu w} \hat{m}_{t+j-1} \right]
 \end{aligned} \tag{1.93}$$

in which $\beta \equiv \pi\mu/R$ is the inverse of the steady state growth-adjusted real interest rate and $\hat{\Delta}_t = \sum_{k=1}^j (\hat{\pi}_{t+k} + \hat{\mu}_{t+k} - \hat{R}_{t+k-1})$. Simplifying (1.93) further and expressing variables in logarithm, I finally get

$$\begin{aligned}\tilde{w}_t &= k_2 + E_t \sum_{j=1}^{\infty} \beta^j \left[\frac{\tau}{w} \tilde{\tau}_{t+j} - \frac{g}{w} \tilde{g}_{t+j} + \left(\kappa - \frac{(R-1)m}{\pi\mu w} \right) \tilde{\mu}_{t+j} \right] \\ &+ E_t \sum_{j=1}^{\infty} \beta^j \left[\left(\kappa - \frac{(R-1)m}{\pi\mu w} \right) \tilde{\pi}_{t+j} - \left(\kappa - \frac{Rm}{\pi\mu w} \right) \tilde{R}_{t+j-1} + \frac{(R-1)m}{\pi\mu w} \tilde{m}_{t+j-1} \right]\end{aligned}$$

in which $\kappa = \left[\frac{s}{w} + \frac{(R-1)m}{\pi\mu w} \right] \left(\frac{1}{1-\beta} \right)$ and

$$\begin{aligned}k_2 &\equiv \ln(w) - \frac{\beta}{1-\beta} \left[\kappa \ln(\beta) + \frac{\tau}{w} \ln(\tau) - \frac{g}{w} \ln(g) - \frac{(R-1)m}{\pi\mu w} \ln(\mu) \right] \\ &+ \frac{\beta}{1-\beta} \left[\frac{(R-1)m}{\pi\mu w} \ln(\pi) - \frac{Rm}{\pi\mu w} \ln(R) - \frac{(R-1)m}{\pi\mu w} \ln(m) \right]\end{aligned}$$

Appendix 1.C.2: Deriving (1.51)-(1.52)

First, I apply (1.50) in (1.45) and get

$$\begin{aligned}\bar{f}_t \bar{C}_w &= k_2 + E_t \sum_{j=1}^{\infty} \beta^j \left(\frac{\tau}{w} \bar{f}_{t+j} \bar{C}_\tau - \frac{g}{w} \bar{f}_{t+j} \bar{C}_g + \kappa_1 \bar{f}_{t+j} \bar{C}_\mu + \kappa_1 \bar{f}_{t+j} \bar{C}_\pi \right) \\ &- E_t \sum_{j=1}^{\infty} \beta^j \left(\kappa_2 \bar{f}_{t+j-1} \bar{C}_R - \kappa_3 \bar{f}_{t+j-1} \bar{C}_m \right)\end{aligned}\tag{1.94}$$

in which $\kappa_1 = \kappa - \frac{(R-1)m}{\pi\mu w}$, $\kappa_2 = \kappa - \frac{Rm}{\pi\mu w}$, $\kappa_3 = \frac{(R-1)m}{\pi\mu w}$.

According to (1.49), I have

$$E_t \bar{f}_{t+j} = \bar{f}_t B^j + \bar{B}_0 \sum_{k=0}^{j-1} B^k \quad (1.95)$$

Applying (1.95) in (1.94) and rearranging terms, I get

$$\begin{aligned} \bar{f}_t \bar{C}_w &= k_2 + \sum_{j=1}^{\infty} \beta^j \left[\left(\bar{f}_t B^j + \bar{B}_0 \sum_{k=0}^{j-1} B^k \right) \left(\frac{\tau}{w} \bar{C}_\tau - \frac{g}{w} \bar{C}_g + \kappa_1 \bar{C}_\mu + \kappa_1 \bar{C}_\pi \right) \right] \\ &\quad - \sum_{j=1}^{\infty} \beta^j \left[\left(\bar{f}_t B^j + \bar{B}_0 \sum_{k=0}^{j-1} B^k \right) (\kappa_2 \beta \bar{C}_R - \kappa_3 \beta \bar{C}_m) \right] \\ &\quad - \kappa_2 \beta \bar{f}_t \bar{C}_R + \kappa_3 \beta \bar{f}_t \bar{C}_m \end{aligned} \quad (1.96)$$

Assuming the VAR(1) process (1.49) is stationary, it can be shown that

$$\sum_{j=1}^{\infty} \beta^j \sum_{k=0}^{j-1} B^k = \frac{\beta}{1-\beta} \sum_{j=1}^{\infty} (\beta B)^{j-1} = \frac{\beta}{1-\beta} (\mathbf{I} - \beta B)^{-1} \quad (1.97)$$

Applying (1.97) in (1.96), I get

$$\begin{aligned} &\left[\beta \bar{f}_t (\mathbf{I} - \beta B)^{-1} B + \frac{\beta}{1-\beta} \bar{B}_0 (\mathbf{I} - \beta B)^{-1} \right] * \\ &\left(\frac{\tau}{w} \bar{C}_\tau - \frac{g}{w} \bar{C}_g + \kappa_1 \bar{C}_\mu + \kappa_1 \bar{C}_\pi - \kappa_2 \beta \bar{C}_R + \kappa_3 \beta \bar{C}_m \right) \\ &+ k_2 = \bar{f}_t (\kappa_2 \beta \bar{C}_R - \kappa_3 \beta \bar{C}_m + \bar{C}_w) \end{aligned} \quad (1.98)$$

Collecting terms of (1.98) that are with \bar{f}_t and rearranging, I obtain the following restriction on B .

$$\beta B \left\{ \frac{\tau}{w} \bar{C}_\tau - \frac{g}{w} \bar{C}_g + \kappa_1 \bar{C}_\mu + \kappa_1 \bar{C}_\pi + \bar{C}_w \right\} = \left(\kappa \beta - \frac{m}{w} \right) \bar{C}_R - \frac{(R-1)m}{Rw} \bar{C}_m + \bar{C}_w$$

Collecting terms of (1.98) that are without \bar{f}_t and rearranging, I obtain the following restriction on \bar{B}_0 .

$$\begin{aligned} & k_2 + \left(\frac{\beta}{1-\beta} \right) \bar{B}_0 (I - \beta B)^{-1} \left(\frac{\tau}{w} \bar{C}_\tau - \frac{g}{w} \bar{C}_g + \kappa_1 \bar{C}_\mu + \kappa_1 \bar{C}_\pi \right) \\ & + \left(\frac{\beta}{1-\beta} \right) \bar{B}_0 (I - \beta B)^{-1} \left[\left(\frac{m}{w} - \kappa\beta \right) \bar{C}_R + \frac{(R-1)m}{Rw} \bar{C}_m \right] = 0 \end{aligned}$$

Appendix 1.C.3: The FGLS Estimator

As shown in Chung and Leeper (2009) and Traum (2007), the part of the objective function (1.57) that is related to b can be rewritten as $-(b - \hat{b})' S (b - \hat{b})$, in which $\hat{b} = \text{vec}((X'^{-1}(X'F))$ and $S \equiv (\Sigma^{-1} \otimes X'X)$. Maximizing (1.57) subject to (1.60), the first-order condition can be shown as $\tilde{b} = \hat{b} - S^{-1}V'\xi$, in which ξ is the Lagrangian multiplier of the constrained optimization problem. From the first-order condition, I can derive $\xi = (VS^{-1}V'^{-1}(V\tilde{b} - V\hat{b}))$. Substituting this expression for ξ back to the first-order condition, I get

$$\tilde{b} = \hat{b} + S^{-1}V'^{-1}V'^{-1}(\text{vec}(\bar{C}_0) - V\hat{b})$$

in which I impose $V\tilde{b} = \text{vec}(\bar{C}_0)$.

Appendix 1.C.4: Constructing the Data Set

I use the quarterly U.S. data from 1947:2 to 2006:2 for the estimation. The data

set is constructed by following Chung and Leeper (2009) and Traum (2007). If not otherwise noted, the data are from the Bureau of Economic Analysis' NIPA tables. All the quantity variables in levels are nominal values, which are divided by the GDP deflator to form real values. For the VAR, I scale all the real quantity variables by the real GDP.

- Y_t : The nominal GDP that is seasonally adjusted at annual rates (line 1 in NIPA Table 1.1.5).
- P_t : The GDP deflator for personal consumption expenditures that is seasonally adjusted (line 2 in NIPA Table 1.1.4).
- y_t : The real GDP, i.e. Y_t/P_t .
- μ_t : The real GDP growth rate, i.e. y_t/y_{t-1} .
- π_t : The GDP deflator inflation rate, i.e. P_t/P_{t-1} .
- R_t : The (annualized) three-month Treasury bill secondary market rate (Source: Board of Governors of the Federal Reserve System, Statistics & Historical Data). Quarterly rates are constructed by converting the annualized rates (monthly frequency) to quarterly rates (monthly frequency) and then taking averages of the three quarterly rates (monthly frequency) within each quarter.
- M_t : The nominal St. Louis adjusted monetary base that is seasonally adjusted (Source: Federal Reserve Bank of St. Louis, Economic Data-FRED, series AMBSL). Quarterly data are constructed by taking averages of monthly data.

- m_t : The real monetary base, i.e. M_t/P_t .
- T_t : The nominal federal taxes that are defined as the sum of all current tax receipts (line 2 in NIPA Table 3.2) and contributions for social insurance (line 11 in NIPA Table 3.2). The original data are seasonally adjusted at annual rates.
- G_t : The nominal federal spending that is defined as the sum of federal consumption expenditure (line 20 in NIPA Table 3.2), gross government investment (line 42 in NIPA Table 3.2) and net purchases of nonproduced assets (line 44 in NIPA Table 3.2), minus consumption of fixed capital (line 45 in NIPA Table 3.2). The original data are seasonally adjusted at annual rates.
- Z_t : The nominal net transfers that are defined as the sum of net current transfers, net capital transfers and subsidies (line 32 in NIPA Table 3.2), minus income receipts on assets (line 12 in NIPA Table 3.2) and current surplus of government enterprises (line 19 in NIPA Table 3.2). Net current transfers are defined as current transfer payments (line 22 in NIPA Table 3.2) minus current transfer receipts (line 16 in NIPA Table 3.2). Net capital transfers are defined as capital transfer payments (line 43 in NIPA Table 3.2) minus capital transfer receipts (line 39 in NIPA Table 3.2). The original data are seasonally adjusted at annual rates.
- τ_t : The real net taxes, i.e. $T_t/P_t - Z_t/P_t$.
- g_t : The real federal spending, i.e. G_t/P_t .

- V_t : The nominal federal debt that is constructed to obey the following government budget constraint

$$V_t - V_{t-1} = \text{Net Borrowing} - \text{Seigniorage}$$

in which the nominal seigniorage is defined as $M_t - M_{t-1}$. The nominal net borrowing is defined as $G_t + INT_t + Z_t - T_t$, where INT_t is interest payments (line 29 in NIPA Table 3.2). To construct the V_t series from 1947:2, I set the value of 1947:1 debt according to the Cox and Hirschhorn (1983) data set³¹. To guarantee the validity of the constructed debt series, I compare it with the Cox-Hirschhorn series and they turn out to be consistent.

- w_t : The real total government liabilities, i.e. $V_t/P_t + M_t/P_t$.

Appendix 1.C.5: Constructing Elasticities for the Identified VAR

- The elasticity of net taxes/GDP to real GDP growth rate, i.e. $\alpha_{\tau\mu}$:

Since $\tau_t/y_t = t_t/y_t - z_t/y_t$, applying log linearization around the stationary steady state, I get

$$\alpha_{\tau\mu} = \frac{\widetilde{d(\tau_t/y_t)}}{\widetilde{d(y_t/y_{t-1})}} = \frac{t/y}{\tau/y} \frac{\widetilde{d(t_t/y_t)}}{\widetilde{d(y_t/y_{t-1})}} - \frac{z/y}{\tau/y} \frac{\widetilde{d(z_t/y_t)}}{\widetilde{d(y_t/y_{t-1})}}$$

³¹The Cox-Hirschhorn debt series is available at <http://www.dallasfed.org/data/data/natdebt.htm>. Following Chung and Leeper (2009) and Traum (2007), I choose not to use the Cox-Hirschhorn debt series because the series is not consistent with NIPA's definition of net borrowing.

in which the steady-state ratios t/y , z/y and τ/y are calibrated as 0.0442, 0.0213 and 0.0229, respectively. It can also be shown that

$$\begin{aligned}\frac{\widetilde{d(t_t/y_t)}}{\widetilde{d(y_t/y_{t-1})}} &= \frac{d\tilde{t}_t - d\tilde{y}_t}{d\tilde{y}_t - d\tilde{y}_{t-1}} = \frac{d\tilde{t}_t}{d\tilde{y}_t} - 1 \\ \frac{\widetilde{d(z_t/y_t)}}{\widetilde{d(y_t/y_{t-1})}} &= \frac{d\tilde{z}_t - d\tilde{y}_t}{d\tilde{y}_t - d\tilde{y}_{t-1}} = \frac{d\tilde{z}_t}{d\tilde{y}_t} - 1\end{aligned}$$

in which $d\tilde{y}_{t-1} = 0$ because the information at time $t-1$ is taken as given. For $d\tilde{t}_t/d\tilde{y}_t$ and $d\tilde{z}_t/d\tilde{y}_t$, I take the values calibrated in Chung and Leeper (2009), which are 3.15 and -0.15, respectively. Therefore, I calibrate $\alpha_{\tau\mu}$ to be 5.2146.

- The elasticity of net taxes/GDP to inflation rate, i.e. $\alpha_{\tau\pi}$:

Similarly to the case of $\alpha_{\tau\mu}$, I have

$$\alpha_{\tau\pi} = \frac{\widetilde{d(\tau_t/y_t)}}{d\tilde{\pi}_t} = \frac{t/y}{\tau/y} \frac{\widetilde{d(t_t/y_t)}}{d\tilde{\pi}_t} - \frac{z/y}{\tau/y} \frac{\widetilde{d(z_t/y_t)}}{d\tilde{\pi}_t}$$

Approximately, I have

$$\begin{aligned}\frac{\widetilde{d(t_t/y_t)}}{d\tilde{\pi}_t} &\approx \frac{d\tilde{t}_t}{d\tilde{\pi}_t} \\ \frac{\widetilde{d(z_t/y_t)}}{d\tilde{\pi}_t} &\approx \frac{d\tilde{z}_t}{d\tilde{\pi}_t}\end{aligned}$$

For $d\tilde{t}_t/d\tilde{\pi}_t$ and $d\tilde{z}_t/d\tilde{\pi}_t$, I take the values calibrated in Chung and Leeper (2009), which are 1.64 and -1, respectively. Therefore, I calibrate $\alpha_{\tau\pi}$ to be 4.0917.

- The elasticity of net taxes/GDP to interest rate, i.e. $\alpha_{\tau R}$:

Similarly as above, I have

$$\alpha_{\tau R} = \frac{d(\widetilde{\tau_t/y_t})}{d\tilde{R}_t} = \frac{t/y}{\tau/y} \frac{d(\widetilde{t_t/y_t})}{d\tilde{R}_t} - \frac{z/y}{\tau/y} \frac{d(\widetilde{z_t/y_t})}{d\tilde{R}_t}$$

Approximately, I have

$$\begin{aligned} \frac{d(\widetilde{t_t/y_t})}{d\tilde{R}_t} &\approx \frac{d\tilde{t}_t}{d\tilde{R}_t} \\ \frac{d(\widetilde{z_t/y_t})}{d\tilde{R}_t} &\approx \frac{d\tilde{z}_t}{d\tilde{R}_t} \end{aligned}$$

Following Perotti (2004) and Chung and Leeper (2009), I calibrate both $d\tilde{t}_t/d\tilde{R}_t$ and $d\tilde{z}_t/d\tilde{R}_t$ to be 0. Therefore, I calibrate $\alpha_{\tau R}$ to be 0.

- The elasticity of government spending/GDP to real GDP growth rate, i.e. $\alpha_{g\mu}$:

Obviously,

$$\alpha_{g\mu} = \frac{d(\widetilde{g_t/y_t})}{d(\widetilde{y_t/y_{t-1}})} = \frac{d\tilde{g}_t - d\tilde{y}_t}{d\tilde{y}_t - d\tilde{y}_{t-1}} = \frac{d\tilde{g}_t}{d\tilde{y}_t} - 1$$

where $d\tilde{y}_{t-1} = 0$. Following Chung and Leeper (2009), I calibrate $d\tilde{g}_t/d\tilde{y}_t$ to be 0. Therefore, $\alpha_{g\mu}$ is calibrated to be -1.

- The elasticity of government spending/GDP to inflation rate, i.e. $\alpha_{g\pi}$:

Approximately,

$$\alpha_{g\pi} = \frac{d(\widetilde{g_t/y_t})}{d\tilde{\pi}_t} \approx \frac{d\tilde{g}_t}{d\tilde{\pi}_t}$$

Following Chung and Leeper (2009), I calibrate $\alpha_{g\pi}$ to be -0.5.

- The elasticity of government spending/GDP to interest rate, i.e. α_{gR} :

Approximately,

$$\alpha_{gR} = \frac{d(\widetilde{g_t/y_t})}{d\tilde{R}_t} \approx \frac{d\tilde{g}_t}{d\tilde{R}_t}$$

Following Perotti (2004) and Chung and Leeper (2009), I calibrate α_{gR} to be 0.

GMM Estimation	$\hat{\theta}_{GMM}$	t-stat	p-value	J-stat	Prob($\chi^2(q) > \text{J-stat}$)
Case 1 (1952:1-2006:2)	0.0492	3.29	0.0010	4.6737	0.1973
Case 2 (1959:1-2005:4)	0.0674	4.06	0.0000	7.0770	0.1319
Case 3 (1960:1-2004:4)	0.0685	4.24	0.0000	10.5187	0.0618
Case 4 (1971:1-2001:4)	0.0763	3.51	0.0005	10.1615	0.1180

Table 1.1: Baseline GMM estimate of (1.4).

Note: Only the result of $\hat{\theta}_{GMM}$ is reported. The default null hypothesis is $H_0 : \theta = 0$. Prob($\chi^2(q) > \text{J-stat}$) is the p-value associated with J-stat, which is the J-statistic of Hansen's test of overidentifying restrictions. Data are at quarterly frequency. Time span for each case is listed in the first column. In case 1, $\tilde{z}_t^1 = \{1, \tau_{t-1}, x_{t-1}, \pi_{t-1}, i_{t-1}, b_{t-1}, PPI_{t-1}, i_{t-1}^{Long}, NAIRU_{t-1}\}$. In case 2, $\tilde{z}_t^2 = \{\tilde{z}_t^1, M3_{t-1}\}$. In case 3, $\tilde{z}_t^3 = \{\tilde{z}_t^2, i_{t-1}^{Germany}\}$. In case 4, $\tilde{z}_t^4 = \{\tilde{z}_t^3, DEMUSD_{t-1}\}$.

GMM Estimation	$\hat{\theta}_{GMM}$	t-stat	p-value	J-stat	Prob($\chi^2(q) > \text{J-stat}$)
Case 1 (1952:1-2006:2)	0.0535	3.72	0.0002	12.5987	0.3204
Case 2 (1959:1-2005:4)	0.0558	4.73	0.0000	19.1550	0.1184
Case 3 (1960:1-2004:4)	0.0606	4.61	0.0000	19.9713	0.1730
Case 4 (1971:1-2001:4)	0.0521	5.49	0.0000	15.0227	0.5938

Table 1.2: GMM estimate of (1.4) with expanded set of instrumental variables.

GMM Estimation	$\hat{\theta}_{GMM}$	t-stat	p-value	J-stat	Prob($\chi^2(q) > \text{J-stat}$)
Case 1 (1952-2006)	0.1814	3.96	0.0001	1.0771	0.7826
Case 2 (1959-2005)	0.3024	5.54	0.0000	5.6154	0.2298
Case 3 (1960-2004)	0.3038	6.16	0.0000	5.7817	0.3280
Case 4 (1971-2001)	0.3364	8.00	0.0000	4.1757	0.6529

Table 1.3: GMM estimate of (1.4) with annual data.

Present-Value Component	Net Taxes Shock	Spending Shock
Net Taxes	-4.3222 [-2.1060,0.5656]	67.6013 [-4.4082,3.4065]
Spending	2.5547 [-0.6101,1.5057]	-42.4831 [-2.5342,3.5157]
Primary Surplus	-1.7674 [-0.9204,0.1510]	25.1182 [-1.2943,1.0413]
Discount Rate	2.1825 [-0.1275,1.1199]	-30.8156 [-1.1358,1.5332]
Seigniorage	-0.4075 [-0.1581,-0.0005]	5.7412 [-0.2581,0.1825]

Table 1.4: Present-value components of fiscal financing (baseline estimation).

Note: Median estimate of each component is provided along with 90% confidence interval. The confidence intervals are estimated with the bias-adjusted bootstrap procedure proposed by Kilian (1998). Fiscal shocks are normalized to raise debt/GDP from the steady state by 1%. The net taxes and spending components sum to the primary surplus component. The primary surplus, discount rate and seigniorage components sum to the 1% increase in debt/GDP.

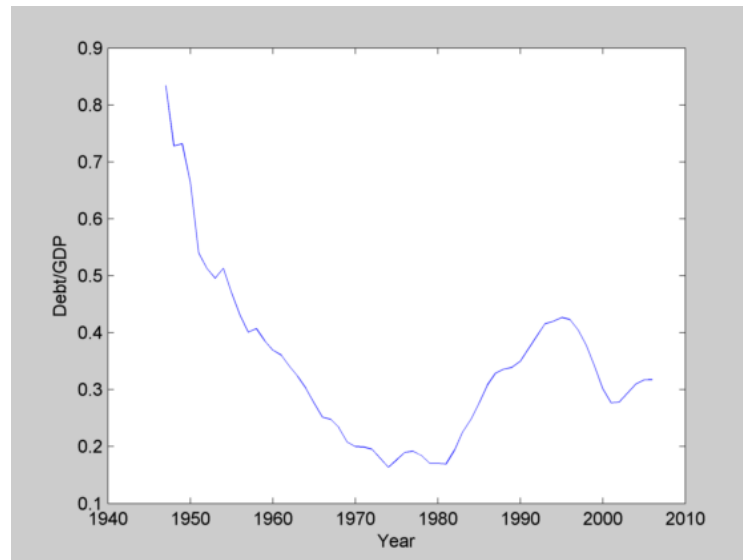


Figure 1.1: U.S. Debt to GDP Ratio (1947-2006, Annually)

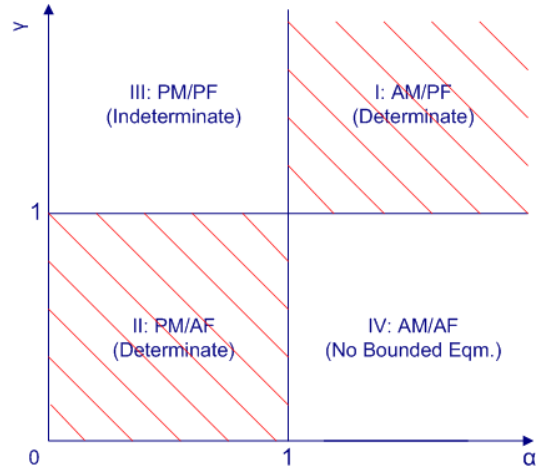
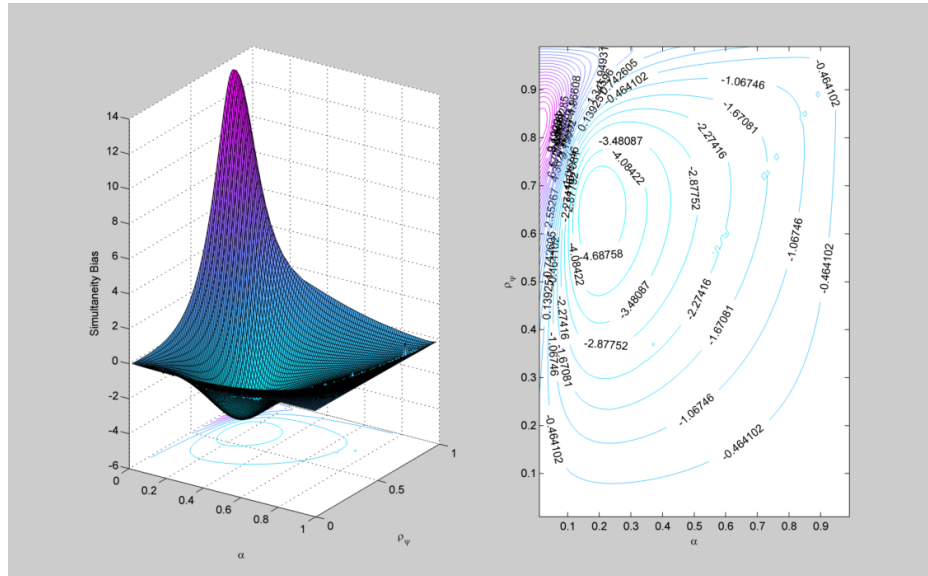
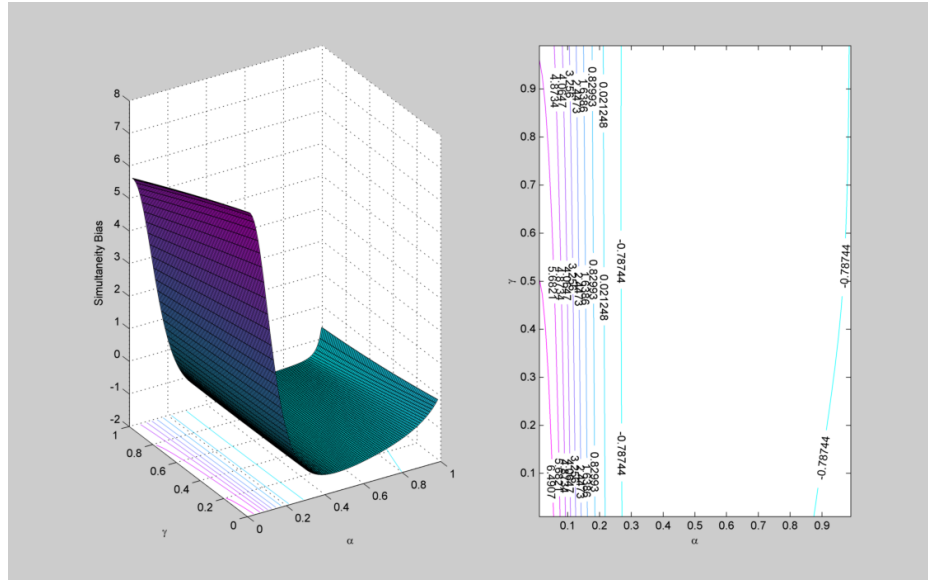


Figure 1.2: First Quadrant of (α, γ) Space





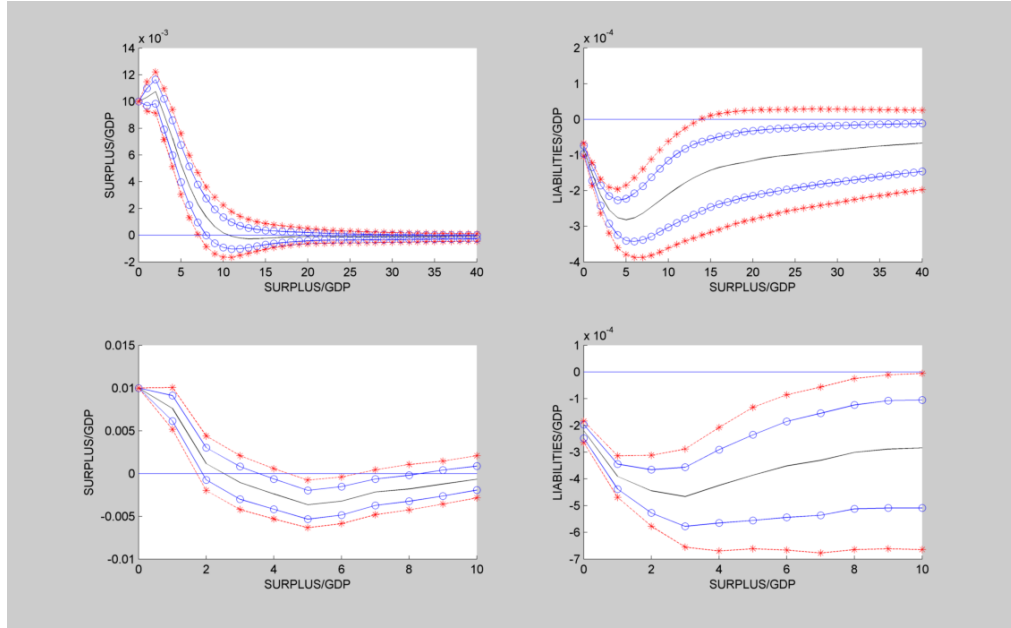


Figure 1.6: Impulse responses of Surplus/GDP and Liabilities/GDP to a positive Surplus/GDP shock.

Note: VAR ordering, Surplus/GDP, Liabilities/GDP. Identification is based on Cholesky decomposition. In the top panel, VAR is estimated with three lags and a constant over the sample 1947:2-2006:2 using the U.S. quarterly data. In the bottom panel, VAR is estimated with five lags and a constant over the sample 1947-2006 using the U.S. annual data. In all the plots, the black solid line represents the median response. The blue solid line with circle represent the bounds of 68% confidence interval. The red dashed line with asterisk represent the bounds of 90% confidence interval. Confidence intervals are estimated with the bias-adjusted bootstrap procedure proposed by Kilian (1998).

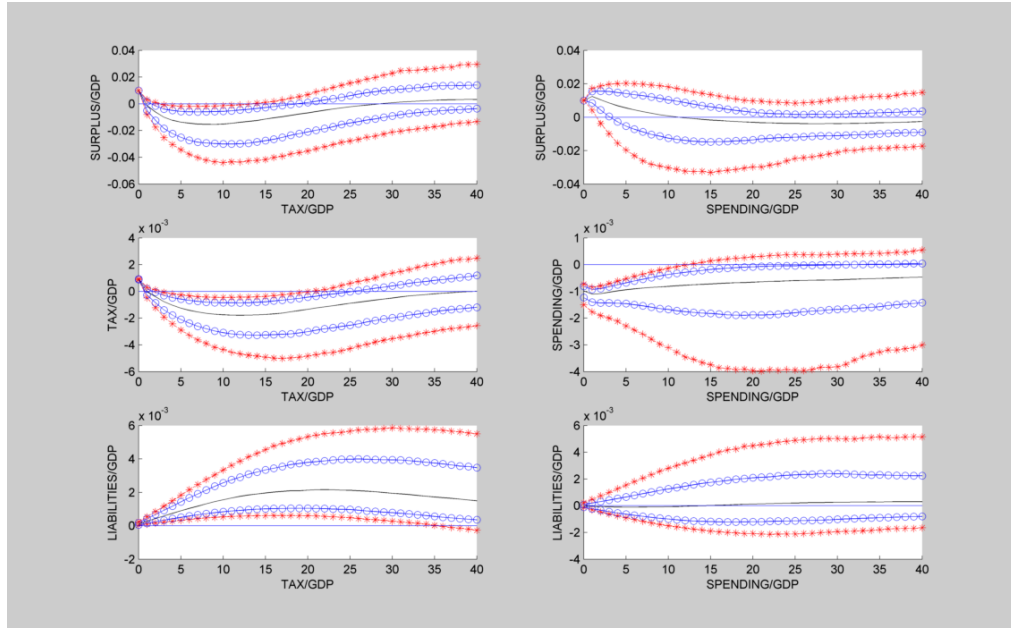


Figure 1.7: Impulse responses of surplus/GDP and liabilities/GDP from the baseline estimation.

Note: Fiscal shocks are normalized to raise surplus/GDP from the steady state by 1%. The left panel is following a surprise net taxes hike. The right panel is following a surprise spending cut. The black solid line represents the median response. The blue solid line with circle represent the bounds of 68% confidence interval. The red dashed line with asterisk represent the bounds of 90% confidence interval. The confidence intervals are estimated with the bias-adjusted bootstrap procedure proposed by Kilian (1998).

Chapter 2

On the Identification of Fiscal Policy Behavior: A Bayesian DSGE Approach

Abstract

The identification of fiscal policy behavior is a general equilibrium problem, which requires that monetary policy behavior always be considered simultaneously. This paper takes a holistic view and jointly identifies monetary and fiscal policy behavior for the U.S. by estimating a standard New-Keynesian sticky-price model with Bayesian methods. By applying Bayesian model comparison techniques to the U.S. Pre-Volcker and Post-1982 samples, this paper finds out that (1) both samples favor determinacy over indeterminacy; (2) active fiscal policy is not detected in either

sample. The findings are consistent with the previous literature for the Post-1982 sample, but not for the Pre-Volcker sample.

2.1 Introduction

It is a widely accepted notion that the nature of the equilibrium of the macro economy hinges on the policy interaction between monetary and fiscal authorities. Therefore, to identify fiscal policy behavior¹ and learn the nature of the economy, monetary policy should always be considered simultaneously. However, the empirical literature conventionally treats monetary and fiscal policy as a dichotomy, i.e. only one kind of policy behavior is studied and the other is usually ignored². In recent years, several people have pointed out that identifying fiscal policy behavior but ignoring monetary policy tends to deliver unreliable inferences³. Woodford (1998) suggests imposing more structures on the model and estimating monetary and fiscal policies simultaneously.

To this end, this paper takes a holistic view and jointly identifies monetary and

¹This paper only focuses on fiscal authority's stance on debt stabilization. The other aspects of fiscal policy behavior, such as fiscal authority's stance on economic stabilization, are ignored for simplicity.

²Typical methods to identify fiscal authority's stance on debt stabilization but ignoring monetary policy behavior include running Ordinary Least Squares (OLS) regression on a fiscal policy rule, e.g. Bohn (1998), running Generalized Method of Moments (GMM) estimation on a fiscal policy rule, e.g. Ballabriga and Martinez-Mongay (2003) and Claeys (2006, 2008), estimating a bivariate fiscal Vector Autoregression (VAR), e.g. Canzoneri, Cumby and Diba (2001, 2002) and Creel and Bihan (2006), and estimating a dynamic stochastic general equilibrium (DSGE) model, e.g. Leeper, Plante and Traum (2010). Typical methods to identify monetary policy behavior but ignoring fiscal policy behavior include running GMM estimation on a monetary policy rule, e.g. Clarida, Galí and Gertler (2000), estimating a monetary VAR, e.g. Christiano, Eichenbaum and Evans (1998), and estimating a DSGE model, e.g. Lubik and Schorfheide (2004).

³Typical examples include Cochrane (1998), Woodford (1998, 2001), Davig, Leeper and Chung (2007) and Li (2010a), among others.

fiscal policy behavior for the U.S. by estimating a standard New-Keynesian sticky-price model with Bayesian methods. In the recently growing literature of estimated new-Keynesian models, the U.S. monetary policy is usually estimated to be passive⁴ in the Pre-Volcker sample and active in the Post-1982 sample⁵. But fiscal policy is usually trivialized by the assumption that lump-sum taxes/transfers adjust to clear the government budget constraint each period. In another word, fiscal policy is always assumed to be passive. Thus, the equilibrium of the Pre-Volcker sample in the U.S. is generally inferred as being indeterminate and that of the Post-1982 sample is inferred as determinate, as concluded in Lubik and Schorfheide (2004). A potential pitfall of this conclusion is that fiscal policy, which should be identified jointly with monetary policy, is trivialized by assumption and active fiscal policy is ruled out a priori. As a paper moving one step forward, Caivano (2007) allows for active fiscal policy and estimates a small-scale new-Keynesian model similar to Woodford (1996) with Bayesian methods. According to Caivano (2007), indeterminacy is a more plausible interpretation of the U.S. Pre-Volcker sample than a non-Ricardian one, which means that the U.S. fiscal policy in the Pre-Volcker sample was passive, which is consistent with the widely held assumption. But Caivano (2007) does not provide a formal test of indeterminacy as in Lubik and Schorfheide (2004). This paper synthesizes the previous work by adopting Lubik and Schorfheide (2004)'s formal

⁴According to Leeper's (1991) terminology, monetary policy is passive when the interest rate responds to inflation at a value of less than one-for-one and active when the response is at a value of more than one-for-one. Fiscal policy is passive when the tax responds to the outstanding government debt strongly and active otherwise. According to Woodford's (1995) terminology, an equilibrium is called non-Ricardian when the monetary policy is passive and the fiscal policy is active. Symmetrically, an equilibrium is called Ricardian when the monetary policy is active and the fiscal policy is passive.

⁵According to Lubik and Schorfheide (2004), the Pre-Volcker sample refers to 1960:I to 1979:II and the Post-1982 sample refers to 1982:IV to 1997:IV.

test of indeterminacy and applying it to the estimated model that allows for active fiscal policy. Taking different policy combinations as alternative data-generating processes, I apply Bayesian model comparison techniques to draw inferences on the policy regimes of the U.S. Pre-Volcker and Post-1982 samples. Since fiscal policy behavior is identified by estimating the whole model with Bayesian methods, which is a full-information approach, the simultaneity bias problem studied in Li (2010a) is avoided a priori.

The main empirical finding of this paper is that both the Pre-Volcker and the Post-1982 samples favor a data-generating process that is characterized by a combination of active monetary policy and passive fiscal policy, which indicates determinacy. The finding is consistent with the previous literature for the Post-1982 sample. For the Pre-Volcker sample, the finding of this paper is at odds with Lubik and Schorfheide (2004) in that an active monetary policy is more favorable than a passive one. The main finding implies that active fiscal policy is not detected in either sample.

2.2 The Model

For the empirical analysis, I use a standard New-Keynesian sticky-price model similar to Woodford (1996), Kim (2003), Davig and Leeper (2006) and Caivano (2007). In the model, there is a utility-maximizing representative household and a continuum of monopolistically competitive firms facing price stickiness that is governed by Calvo (1983) pricing. The government sets fiscal policy according to a tax rule and

the central bank sets monetary policy according to a Taylor-type nominal interest rate rule. There are four types of structural shocks: technology shocks, monetary policy shocks, government spending shocks and tax shocks. Even though the model is presented in great details in the aforementioned papers, I briefly outline the necessary model setup as follows for clarity of the paper.

2.2.1 Model Setup

The representative household chooses a composite consumption good (C_t), labor input (N_t), nominal money balance (M_t) and one-period nominal government bond (B_t) to maximize

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{C_t^{1-\sigma}}{1-\sigma} + \delta \frac{(M_t/P_t)^{1-\kappa}}{1-\kappa} - \chi \frac{N_t^{1+\nu}}{1+\nu} \right] \quad (2.1)$$

with $0 < \beta < 1$, $\sigma > 0$, $\kappa > 0$, $\nu > 0$, $\delta > 0$ and $\chi > 0$.

The composite consumption good (C_t) is produced by a Dixit and Stiglitz (1977) aggregator:

$$C_t = \left[\int_0^1 c_{jt}^{\frac{\theta-1}{\theta}} dj \right]^{\frac{\theta}{\theta-1}} \quad (2.2)$$

where c_{jt} is demand for differentiated good j and $\theta > 1$ is price elasticity of demand for good j . From the household expenditure minimization problem, it can be shown that

$$c_{jt} = \left(\frac{p_{jt}}{P_t} \right)^{-\theta} C_t \quad (2.3)$$

where p_{jt} is price of the differentiated good j and P_t is aggregate price index for the composite consumption good that is defined as

$$P_t \equiv \left[\int_0^1 p_{jt}^{1-\theta} dj \right]^{\frac{1}{1-\theta}} \quad (2.4)$$

The household budget constraint is

$$C_t + \frac{M_t}{P_t} + \frac{B_t}{P_t} + \tau_t \leq \left(\frac{W_t}{P_t} \right) N_t + \frac{M_{t-1}}{P_t} + \frac{R_{t-1}B_{t-1}}{P_t} + \Pi_t \quad (2.5)$$

where τ_t is lump-sum taxes (if positive) or transfers (if negative), W_t is nominal wage, R_t is nominal interest rate and Π_t is profits from firms that is owned by the household. The household problem is to maximize (2.1) subject to (2.5), which results in the following first-order necessary conditions

$$\chi \frac{N_t^\nu}{C_t^{-\sigma}} = \frac{W_t}{P_t} \quad (2.6)$$

$$R_t^{-1} = \beta E_t \left(\frac{C_t}{C_{t+1}} \right)^\sigma \frac{P_t}{P_{t+1}} \quad (2.7)$$

$$\frac{M_t}{P_t} = \delta^{1/\kappa} \left(\frac{i_t}{1+i_t} \right)^{-1/\kappa} C_t^{\sigma/\kappa} \quad (2.8)$$

where $i_t \equiv R_t - 1$ is net nominal interest rate. Besides (2.6)-(2.8), the following transversality condition must also hold all the time in equilibrium

$$\lim_{T \rightarrow \infty} E_t \left[q_{t,T} \frac{A_T}{P_T} \right] = 0 \quad (2.9)$$

where $A_t \equiv B_t + M_t$ is total nominal wealth and $q_{t,t+1} \equiv (1/R_t)(P_{t+1}/P_t)$ is real discount factor.

The government demands goods in the same way as the households do, which is also based on a Dixit and Stiglitz (1977) aggregator:

$$G_t = \left[\int_0^1 g_{jt}^{\frac{\theta-1}{\theta}} dj \right]^{\frac{\theta}{\theta-1}} \quad (2.10)$$

where g_{jt} , demand for differentiated good j , is

$$g_{jt} = \left(\frac{p_{jt}}{P_t} \right)^{-\theta} G_t \quad (2.11)$$

As one of the standard assumptions of New-Keynesian models, a continuum of monopolistically competitive firms produce differentiated goods using linear technology, which takes labor as input. The production function is

$$y_{jt} = Z_t N_{jt} \quad (2.12)$$

where Z_t is aggregate technology that is common across firms. In this paper, Z_t (in logarithm) is assumed to follow an AR(1) process

$$\log(Z_t) = \rho_Z \log(Z_{t-1}) + \epsilon_t^Z \quad (2.13)$$

where $\rho_Z \in [0, 1)$ and ϵ_t^Z is *i.i.d.* $N(0, \sigma_Z^2)$. From (2.3) and (2.11), the aggregate

demand curve faced by firm j is

$$y_{jt} = \left(\frac{p_{jt}}{P_t} \right)^{-\theta} Y_t \quad (2.14)$$

where $Y_t = C_t + G_t$ is the aggregate resource constraint. Obviously, equating supply and demand for individual goods implies that

$$Z_t N_{jt} = \left(\frac{p_{jt}}{P_t} \right)^{-\theta} Y_t \quad (2.15)$$

Assuming that labor services for individual firms convert into aggregate labor linearly, i.e. $N_t = \int_0^1 N_{jt} dj$, (2.15) implies the aggregate production function as

$$Y_t = \frac{Z_t}{\Delta_t} N_t \quad (2.16)$$

where $\Delta_t \equiv \int_0^1 \left(\frac{p_{jt}}{P_t} \right)^{-\theta} dj$ is a measure of relative price dispersion.

Following Calvo (1983), this paper introduces price stickiness by assuming that a constant fraction $1 - \vartheta$ firms are allowed to re-optimize their prices each period, while the remaining firms are not allowed to adjust. The firms being able to re-optimize their prices actually maximize the present value of their expected future real profit streams, i.e.

$$\max E_t \sum_{s=0}^{\infty} \vartheta^s q_{t,t+s} \Pi_{jt+s} \quad (2.17)$$

where the real profit stream of firm j at period t is

$$\Pi_{jt} = \left(\frac{p_{jt}}{P_t} \right)^{1-\theta} Y_t - \Psi_t \left(\frac{p_{jt}}{P_t} \right)^{-\theta} Y_t \quad (2.18)$$

where Ψ_t , the real marginal cost, is defined in the individual firm's cost minimization problem as

$$\Psi_t = \frac{W_t}{Z_t P_t} \quad (2.19)$$

The first-order condition of the individual firm's optimal pricing problem implies that

$$\frac{p_{jt}^*}{P_t} = \left(\frac{\theta}{\theta - 1} \right) \frac{E_t \sum_{s=0}^{\infty} (\vartheta \beta)^s (Y_{t+s} - G_{t+s})^{-\sigma} \left(\frac{P_{t+s}}{P_t} \right)^{\theta} \Psi_{t+s} Y_{t+s}}{E_t \sum_{s=0}^{\infty} (\vartheta \beta)^s (Y_{t+s} - G_{t+s})^{-\sigma} \left(\frac{P_{t+s}}{P_t} \right)^{\theta-1} Y_{t+s}} \quad (2.20)$$

where p_{jt}^* is the optimal price chosen by firm j and $\frac{\theta}{\theta-1}$ is the markup. Since all the firms being able to re-optimize their prices face the same problem, they will choose the same optimal price in equilibrium, i.e. symmetric monopolistic competitive equilibrium. Therefore, the subscript j of p_{jt}^* will be dropped hereinafter. From (2.4), the aggregate price index evolves according to

$$P_t = [(1 - \vartheta)(p_t^*)^{1-\theta} + \vartheta P_{t-1}^{1-\theta}]^{\frac{1}{1-\theta}} \quad (2.21)$$

Accordingly, the inflation rate evolves according to

$$\pi_t = [(1 - \vartheta)(\pi_t^*)^{1-\theta} + \vartheta \pi^{1-\theta}]^{\frac{1}{1-\theta}} \quad (2.22)$$

where $\pi_t^* \equiv p_t^*/P_{t-1}$. For simplicity, the steady-state inflation rate, i.e. π , is assumed to be 1.

The government sets fiscal policy according to a tax rule, i.e.

$$\tau_t = \gamma_0 + \gamma_b \frac{B_{t-1}}{P_{t-1}} + \tau \psi_t \quad (2.23)$$

where τ is the steady-state value of tax and ψ_t is an exogenous fiscal policy shock following an AR(1) process

$$\psi_t = \rho_\psi \psi_{t-1} + \epsilon_t^\psi \quad (2.24)$$

where ϵ_t^ψ is *i.i.d.* $N(0, \sigma_\psi^2)$. Similarly, the central bank sets monetary policy according to a Taylor-type nominal interest rate rule, i.e.

$$R_t = e^{\alpha_0} \pi_t^{\alpha_\pi} \phi_t \quad (2.25)$$

where ϕ_t is an exogenous monetary policy shock whose logarithm follows an AR(1) process

$$\log(\phi_t) = \rho_\phi \log(\phi_{t-1}) + \epsilon_t^\phi \quad (2.26)$$

where ϵ_t^ϕ is *i.i.d.* $N(0, \sigma_\phi^2)$.

The government budget constraint is

$$G_t = \tau_t + \frac{M_t - M_{t-1}}{P_t} + \frac{B_t}{P_t} - \frac{R_{t-1} B_{t-1}}{P_t} \quad (2.27)$$

In this paper, government spending (in logarithm) is assumed to follow an AR(1) process

$$\log(G_t) = (1 - \rho_G)\log G + \rho_G\log(G_{t-1}) + \epsilon_t^G \quad (2.28)$$

where G is the steady-state value of government spending and ϵ_t^G , i.e. the government spending shock, is *i.i.d.* $N(0, \sigma_G^2)$.

2.2.2 The Linearized Model

The equilibrium conditions of the model are log-linearized around the steady state, which are summarized in Appendix 2.A. The system of first-order difference equations can then be cast into the canonical form

$$\Gamma_0 X_t = \Gamma_1 X_{t-1} + \Pi \eta_t + \Psi \epsilon_t \quad (2.29)$$

where $X_t = [\hat{\pi}_t, \hat{R}_t, \hat{Y}_t, \hat{m}_t, \hat{b}_t, \hat{Z}_t, \hat{G}_t, \hat{\phi}_t, \psi_t]'$, $\eta_t = [\eta_t^\pi, \eta_t^Y]'$ and $\epsilon_t = [\epsilon_t^Z, \epsilon_t^G, \epsilon_t^\phi, \epsilon_t^\psi]'$.

Note, η_t is the vector of one-period-ahead endogenous forecasting error, e.g. $\eta_t^\pi \equiv \hat{\pi}_t - E_{t-1}\hat{\pi}_t$.

2.2.3 Determinacy Regions

It can be shown that the determinacy of bounded equilibrium of the linearized system (2.29) hinges on (α_π, γ_b) , i.e. the two parameters in the monetary and fiscal policy rules. Following Leeper (1991), I characterize four different regions

in the first quadrant of (α_π, γ_b) space⁶, as shown in Figure 2.1: (1) When $\alpha_\pi \in (1, +\infty)$ and $\gamma_b \in (\beta^{-1} - 1, \beta^{-1})$, monetary policy is active (AM) and fiscal policy is passive (PF). The resulting equilibrium is determinate; (2) When $\alpha_\pi \in (0, 1)$ and $\gamma_b \in (0, \beta^{-1} - 1)$, monetary policy is passive (PM) and fiscal policy is active (AF). The resulting equilibrium is also determinate; (3) When $\alpha_\pi \in (0, 1)$ and $\gamma_b \in (\beta^{-1} - 1, \beta^{-1})$, monetary policy and fiscal policy are both passive. The resulting bounded equilibrium is indeterminate; (4) When $\alpha_\pi \in (1, +\infty)$ and $\gamma_b \in (0, \beta^{-1} - 1)$, monetary policy and fiscal policy are both active. Bounded equilibrium does not exist. For the rest of the paper, I only focus on the regions in which bounded equilibrium does exist, i.e. regions (1)-(3).

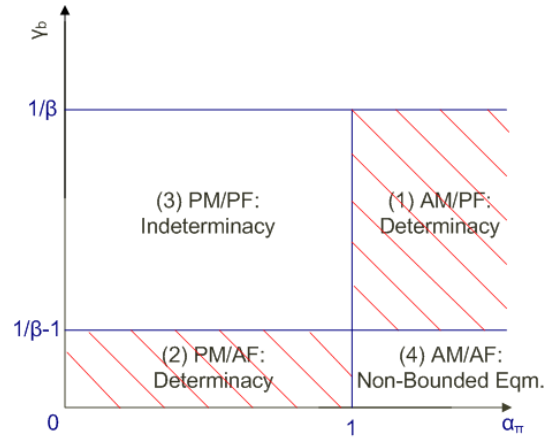


Figure 2.1: First Quadrant of (α_π, γ_b) Space

⁶As to the determinacy regions in the (α_π, γ_b) space, I only focus on those in the first quadrant where both parameters are positive and the associated economic interpretation is meaningful. Moreover, I restrict γ_b to $(0, \beta^{-1})$ so that $\beta^{-1} - \gamma_b$, one of the eigenvalues of the system, is always positive and oscillating dynamics is avoided.

2.2.4 Model Solution

When the bounded equilibrium is determinate, I solve the model using the *gensys* algorithm, which is based on Sims (2001). For the case of indeterminacy, the *gensys* algorithm only provides one particular stable solution, which is not appropriate for the empirical analysis. Therefore, I follow Lubik and Schorfheide (2003, 2004) to compute the full set of stable solutions under indeterminacy. For implementation in practice, I use a modified version of the *gensys* algorithm. For readers' convenience, I briefly sketch the methodology developed in Lubik and Schorfheide (2003) in Appendix 2.B.

In general form, the model solution can be expressed as

$$X_t = G(\Theta)X_{t-1} + K_1(\Theta)\epsilon_t + K_2(\Theta)\zeta_t \quad (2.30)$$

where Θ is the vector of structural parameters to be estimated and ζ_t is the sunspot shock. Since the degree of indeterminacy is at most 1, I set the dimension of ζ_t as 1. Under determinacy, the last term of (2.30) drops out.

2.3 Empirical Analysis

The empirical analysis of this paper is based on the estimation of the model using Bayesian methods. Unlike the recent literature on identifying monetary policy behavior in estimated DSGE models, in which fiscal policy is usually trivialized and

assumed to be passive a priori, e.g. Lubik and Schorfheide (2004), this paper treats fiscal policy as equally important as monetary policy and tries to identify both kinds of policy behavior jointly.

As in Lubik and Schorfheide (2004), this paper will ultimately draw inferences on the policy behavior using Bayesian model comparison techniques. Specifically, I consider the following model specifications according to different combinations of monetary and fiscal policy behavior: \mathcal{M}_1 (AM/PF, i.e. region 1 in Figure 2.1), \mathcal{M}_2 (PM/AF, i.e. region 2) and \mathcal{M}_3 (PM/PF, i.e. region 3).

2.3.1 Estimation Procedure

First, the model is cast into its state-space representation, in which (2.30) serves as the state equation. The observation equation is

$$Q_t = HX_t \quad (2.31)$$

where $Q_t = [\hat{Y}_t^o, \hat{\pi}_t^o, \hat{R}_t^o, \hat{b}_t^o]'$ is the vector of observable variables and H is a selection matrix mapping model variables into observable variables. Note, any observable variable is with a superscript “o”. In this paper, both the observable variables and the model variables are at quarterly frequency.

Based on Bayes’ theorem, the posterior density is defined as

$$p(\Theta|Q_1^T, \mathcal{M}_i) = \frac{\mathcal{L}(\Theta|Q_1^T, \mathcal{M}_i)p(\Theta|\mathcal{M}_i)}{p(Q_1^T|\mathcal{M}_i)}, \quad \text{for } i = \{1, 2, 3\}. \quad (2.32)$$

where $\Theta = [\alpha_\pi, \gamma_b, \sigma, \kappa, \nu, \vartheta, \rho_Z, \rho_G, \rho_\phi, \rho_\psi, \sigma_Z, \sigma_G, \sigma_\phi, \sigma_\psi]$ for \mathcal{M}_1 and \mathcal{M}_2 . For \mathcal{M}_3 , $\Theta = [\alpha_\pi, \gamma_b, \sigma, \kappa, \nu, \vartheta, \rho_Z, \rho_G, \rho_\phi, \rho_\psi, \sigma_Z, \sigma_G, \sigma_\phi, \sigma_\psi, M_1, \sigma_\zeta]$, where M_1 , appearing in (2.45), is related to the indeterminacy of the propagation of fundamental shocks and σ_ζ is the standard deviation of sunspot shock. As in Lubik and Schorfheide (2004), I normalize M_2 in (2.45) to be 1. Q_1^T denotes the whole set of observed data. \mathcal{L} denotes the likelihood function. $p(\Theta|\mathcal{M}_i)$ and $p(Q_1^T|\mathcal{M}_i)$ are the prior density and the marginal data density conditional on \mathcal{M}_i , respectively.

Based on the state-space representation (2.30) and (2.31), the conditional likelihood $\mathcal{L}(\Theta|Q_1^T, \mathcal{M}_i)$ is constructed and evaluated by Kalman filter, as derived in Hamilton (1994). Then the posterior kernel, i.e. $\mathcal{L}(\Theta|Y^o, \mathcal{M}_i)p(\Theta|\mathcal{M}_i)$, is simulated for each model specification by random-walk Metropolis-Hastings (MH) algorithm, which is a Markov Chain Monte Carlo (MCMC) simulation method commonly used in the current literature of Bayesian estimation of DSGE models. To initialize the simulation for each model specification, I use Sims' *csmmwel* routine⁷ to search for the posterior mode and the associated inverse Hessian at the mode in the corresponding region of the parameter space. Then I start the simulation from the posterior mode, treating the inverse Hessian at the mode as the variance-covariance matrix of the proposal density in the random-walk MH algorithm⁸. For a given data set and conditional on each model specification, I run the random-walk MH algorithm for 1 million iterations. Besides the trace plots, I check convergence of the

⁷The *csmmwel* routine is not quite stable near the boundary or cliff in the parameter space. To avoid the non-stability, I start the routine from several over-dispersed points and run the routine in parallel. Then I choose the point with the largest posterior density as the approximated local mode.

⁸Gelman et.al. (2004) and An and Schorfheide (2007) provide explanation of the simulation method in more details.

posterior draws using CODA package and Geweke’s Chi-squared test, according to LeSage (1999) and Geweke (1999)⁹. Acceptance rate is kept between 30% and 40% for different model specifications. After convergence to a stationary distribution is confirmed, I discard the first 10% of the draws as burn-in phase and take every 30th draw of the remaining chain to compute the posterior moments of the parameters.

To draw inferences on policy behavior, I implement Bayesian model comparison and select the most favorable model based on Bayes factor¹⁰. As in Lubik and Schorfheide (2004), An and Schorfheide (2007) and DYNARE (2005), I estimate log marginal data density¹¹, i.e. $\log p(Y^o|\mathcal{M}_i)$, with Geweke’s (1999) modified harmonic-mean (MHM) estimator¹².

2.3.2 Data

The model is estimated with quarterly U.S. data, which include four observable variables: real GDP, inflation rate, nominal interest rate and real government debt. To convert the data into percentage deviations from their steady state levels, I detrend the logarithm of real GDP and real government debt with two independent linear trends. The percentage deviation of inflation from its steady state is easily calculated because its steady state is assumed to be 1. For nominal interest rate,

⁹Some other helpful references on convergence diagnostics are Kass (1998), Gilks et.al. (1996) and Gelman et.al. (2004).

¹⁰Besides Bayes factor, there is another way to draw inferences on competing models, which is to treat model indices as parameters. Simulation is then implemented in the expanded parameter space. A popular example is the reversible jump MCMC algorithm. Justiniano (2004) extensively surveys different strategies for model selection.

¹¹The marginal data density intrinsically penalizes the likelihood function for models with free parameters, such as \mathcal{M}_3 .

¹²Lancaster (2004) also provides a good introduction to the issue of model comparison.

the percentage deviation is calculated relative to its sample mean. Appendix 1.C.4 of Li (2010a) describes the construction of the data in more details.

To be comparable with Lubik and Schorfheide (2004), I consider two sample periods for the estimation¹³: a Pre-Volcker sample from 1960:I to 1979:II and a Post-1982 sample from 1982:IV to 1997:IV.

2.3.3 Prior

Some model parameters are calibrated, as shown in Table 2.1. Specifically, β is calibrated so that the (annualized) steady state real interest rate is 4%. The steady states of the other ratios are their respective sample means.

For the vector of parameters to be estimated, the specified prior densities, means, standard deviations and 90% probability intervals are reported in Table 2.2. It is assumed that all the parameters are independent a priori, which is a common assumption in the literature of Bayesian estimation of DSGE models¹⁴. When a parameter is weakly identified, it is well known that prior may influence the posterior significantly¹⁵. Therefore, this paper takes an agnostic approach on the prior assumptions for the parameters. The monetary policy parameter α_π is centered at 1.1 and the 90% probability interval is from 0.43 to 2.03, which spans a quite large

¹³The Volcker-Greenspan sample from 1979:III to 1997:IV, which is also explored in Lubik and Schorfheide (2004), is not considered in this paper because it has been widely recognized that the monetary policy behavior during the Volcker-disinflation period, i.e 1979:III to 1982:III, cannot be well described by a Taylor-type interest rate rule.

¹⁴This strong assumption is for simplicity. The generalization that parameters are treated as potentially correlated is nontrivial, which is left for future research.

¹⁵Canova and Sala (2009) provides a detailed analysis on the issue of parameter identification in DSGE models.

interval and incorporates both active and passive policy behavior as possibilities. The fiscal policy parameter γ_b is uniformly distributed¹⁶ from 0 to β^{-1} , which incorporates both active and passive policy behavior as possibilities and guarantees that the dynamics is well-behaved. The preference parameters σ , κ and ν all follow Gamma distribution with mean 2 and standard deviation 1. The probability that a firm cannot re-optimize its price follows a Beta distribution. The prior mean 0.66 indicates that on average only one third of the firms are allowed to adjust their prices each period. All of the autoregressive parameters of shock processes follow Beta distribution with mean 0.7 and standard deviation 0.1. Their 90% intervals indicate that the processes are relatively highly persistent. The standard deviations of structural shocks all follow Inverse Gamma distribution with relatively high dispersion. For the parameters characterizing indeterminacy of the propagation of fundamental shocks, i.e. M_1 , uniform priors from -3 to 3 are assumed¹⁷. For \mathcal{M}_3 , two cases are considered. First, there is no sunspot shock. So σ_ζ is not identifiable. Indeterminacy only comes from the propagation of fundamental shocks. Second, there is sunspot shock. So both the fundamental shocks and the sunspot shock contribute to the indeterminacy.

¹⁶Unlike the monetary policy parameter, the literature has not reached a consensus on the prior for the fiscal policy parameter. For example, Caivano (2007) specifies a Beta prior for γ_b , while Leeper, Plante and Traum (2010) specifies a Gamma prior. This paper takes an extremely agnostic approach.

¹⁷Lubik and Schorfheide (2004) treat M_1 in a different way, assuming they are normally distributed with mean at a baseline indeterminacy solution. Specifically, for every parameter vector Θ in the indeterminacy region, Lubik and Schorfheide (2004) construct a corresponding parameter vector $\tilde{\Theta}$ on the boundary of the determinacy region. They then choose M_1^* for the baseline indeterminacy solution by minimizing the discrepancy between the impulse responses conditional on Θ and $\tilde{\Theta}$. This methodology is not directly applicable to this paper because the model in this paper has two possible determinacy regions. To determine which boundary to choose is nontrivial, which is left for future research.

2.3.4 Posterior Inferences

Table 2.3 and 2.4 report the log marginal data density and the Bayes factors conditional on the two different samples. For \mathcal{M}_3 , \mathcal{M}_3^a and \mathcal{M}_3^b denote indeterminacy with and without sunspot shocks, respectively.

For the Pre-Volcker sample, \mathcal{M}_1 is the most favorable data-generating process according to the Bayes factors, which is at odds with Lubik and Schorfheide (2004) where indeterminacy under all different prior assumptions is always more favorable than determinacy. In this paper, not only \mathcal{M}_1 but also \mathcal{M}_2 , which is not considered in Lubik and Schorfheide (2004), is more favorable than \mathcal{M}_3^a , where the latter is corresponding to indeterminacy under Prior 1 in Lubik and Schorfheide (2004). However, compared with \mathcal{M}_3^b that is corresponding to indeterminacy under Prior 3 in Lubik and Schorfheide (2004), \mathcal{M}_2 is less favorable. For the Post-1982 sample, \mathcal{M}_1 is again the most favorable data-generating process, which is consistent with the conclusions in Lubik and Schorfheide (2004). The other findings for the Pre-Volcker sample carry over to this period. In a word, the data-generating process characterized by a combination of active monetary policy and passive fiscal policy is always preferred.

Table 2.5 reports the posterior means and 90% intervals for both sample periods. Since \mathcal{M}_1 is always preferred regardless of the sample, posterior moments in Table 2.5 are all conditional on \mathcal{M}_1 , i.e. the specification with an active monetary policy and a passive fiscal policy. Compared with Lubik and Schorfheide (2004), the magnitude of α_π and σ in this paper are much larger, which may be because a different

data set is used. ρ_Z is almost 1, which implies that the technology process is very likely to have a unit root¹⁸. According to σ_G and σ_ψ , it is clear that the fiscal policy shocks are more volatile than the technology shock and the monetary policy shock.

For readers' reference, trace plots conditional on different model specifications for both periods are displayed in Figure 2.2 to 2.9¹⁹.

2.4 Conclusion

This paper jointly identifies monetary and fiscal policy behavior for the U.S. by estimating a standard New-Keynesian sticky-price model with Bayesian methods. A conventional monetary DSGE model is generalized by allowing for active fiscal policy behavior, which is usually ruled out a priori by the literature of estimated DSGE models. By applying Bayesian model comparison techniques to the U.S. Pre-Volcker and Post-1982 samples, this paper finds out that the combination of an active monetary policy and a passive fiscal policy, which indicates determinacy, is always more favorable than the other policy combinations regardless of the sample. This is consistent with the previous literature for the Post-1982 sample. For the Pre-Volcker sample, this finding is at odds with Lubik and Schorfheide (2004) in that an active monetary policy is more favorable than a passive one. This finding implies that active fiscal policy is not detected in either sample.

¹⁸The issue of unit root, which is not considered in this paper, is left for future work.

¹⁹A challenge lies in the estimation conditional on \mathcal{M}_2 (PM/AF), in which convergence of MCMC is hard to be seen within 1 million draws. Since most of the estimated DSGE models in the current literature assume fiscal policy as passive a priori, this technical challenge deserves further exploration in the future work.

Since the estimated model, data used and prior assumptions in this paper are more or less different from those in Lubik and Schorfheide (2004), plus there are some technical challenges as already mentioned, e.g. convergence problem of MCMC for PM/AF, the empirical findings of this paper need more investigation in the future.

Appendix 2.A: The Linearized Model

Log-linearization of the equilibrium conditions of the model around the steady state yields the following system of equations:

$$\hat{Y}_t = E_t \hat{Y}_{t+1} + \left(\frac{1-g}{\sigma} \right) \left[-\hat{R}_t + E_t \hat{\pi}_{t+1} \right] + g \left[\hat{G}_t - E_t \hat{G}_{t+1} \right] \quad (2.33)$$

$$\begin{aligned} \hat{\pi}_t = & \beta E_t \hat{\pi}_{t+1} + \frac{(1-\vartheta)(1-\vartheta\beta)}{\vartheta} \left(\frac{\sigma}{1-g} + \nu \right) \hat{Y}_t - \frac{(1-\vartheta)(1-\vartheta\beta)(1+\nu)}{\vartheta} \hat{Z}_t \\ & - \frac{(1-\vartheta)(1-\vartheta\beta)}{\vartheta} \left(\frac{\sigma g}{1-g} \right) \hat{G}_t \end{aligned} \quad (2.34)$$

$$\hat{m}_t = \frac{1}{\kappa} \left[\sigma \left(\frac{1}{1-g} \hat{Y}_t - \frac{g}{1-g} \hat{G}_t \right) - \frac{\beta}{1-\beta} \hat{R}_t \right] \quad (2.35)$$

$$\hat{R}_t = \alpha_\pi \hat{\pi}_t + \hat{\phi}_t \quad (2.36)$$

$$\begin{aligned} \left(\frac{m}{b} + \beta^{-1} \right) \hat{\pi}_t + \frac{m}{b} \hat{m}_t + \hat{b}_t = & \beta^{-1} \hat{R}_{t-1} + \frac{m}{b} \hat{m}_{t-1} + (\beta^{-1} - \gamma_b) \hat{b}_{t-1} \\ & + \frac{G}{b} \rho_G \hat{G}_{t-1} - \frac{\tau}{b} \rho_\psi \psi_{t-1} + \frac{G}{b} \epsilon_t^G - \frac{\tau}{b} \epsilon_t^\psi \end{aligned} \quad (2.37)$$

$$\hat{Z}_t = \rho_Z \hat{Z}_{t-1} + \epsilon_t^Z \quad (2.38)$$

$$\hat{G}_t = \rho_G \hat{G}_{t-1} + \epsilon_t^G \quad (2.39)$$

$$\hat{\phi}_t = \rho_\phi \hat{\phi}_{t-1} + \epsilon_t^\phi \quad (2.40)$$

$$\psi_t = \rho_\psi \psi_{t-1} + \epsilon_t^\psi \quad (2.41)$$

where variables without time subscripts denote steady-state values and \hat{x}_t denotes log deviation of the generic variable x_t from its steady-state value x , i.e. $\hat{x}_t \equiv \ln(x_t) - \ln(x)$. Note, $g \equiv G/Y$ is the steady-state ratio of government spending to output. Among the system of equations, (2.33) is the consumption Euler equation; (2.34) is the New-Keynesian Phillips curve; (2.35) is the money demand equation; (2.36) is the linearized monetary policy rule; (2.37) is the linearized government budget constraint; (2.38)-(2.41) are the processes of the structural shocks.

Appendix 2.B: Solution Under Indeterminacy

This appendix highlights the solution strategy under indeterminacy developed in Lubik and Schorfheide (2003), based on which I modify the *gensys* algorithm for estimation.

Following Sims (2001), a generalized complex Schur (QZ) decomposition is applied to Γ_0 and Γ_1 of the canonical form (2.29), i.e. $Q'\Lambda Z' = \Gamma_0$, $Q'\Omega Z' = \Gamma_1$, where $QQ' = Q'Q = ZZ' = Z'Z = I_{n \times n}$ and Λ and Ω are both upper triangular. Premultiplying (2.29) by Q yields

$$\begin{bmatrix} \Lambda_{11} & \Lambda_{12} \\ 0 & \Lambda_{22} \end{bmatrix} \begin{bmatrix} w_{1t} \\ w_{2t} \end{bmatrix} = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ 0 & \Omega_{22} \end{bmatrix} \begin{bmatrix} w_{1t-1} \\ w_{2t-1} \end{bmatrix} + \begin{bmatrix} Q_{1.} \\ Q_{2.} \end{bmatrix} (\Pi\eta_t + \Psi\epsilon_t) \quad (2.42)$$

where $w_t = Z'X_t$ is an $n \times 1$ vector. As stated in Sims (2001), the system can always be ordered and partitioned so that the $m \times 1$ ($0 \leq m \leq n$) vector $w_{2,t}$ is purely explosive. Therefore, the existence of stable solution requires that

$$Q_{2.}\Pi\eta_t + Q_{2.}\Psi\epsilon_t = 0 \quad (2.43)$$

Next, a singular value decomposition is applied to $Q_2.\Pi$, i.e.

$$\begin{aligned}
Q_2.\Pi &= \begin{bmatrix} U_{.1} & U_{.2} \end{bmatrix} \begin{bmatrix} D_{11} & 0 \\ 0 & 0 \end{bmatrix} \begin{bmatrix} V'_{.1} \\ V'_{.2} \end{bmatrix} \\
&= \underbrace{U}_{m \times m} \underbrace{D}_{m \times k} \underbrace{V'}_{k \times k} = \underbrace{U_{.1}}_{m \times r} \underbrace{D_{11}}_{r \times r} \underbrace{V'_{.1}}_{r \times k}
\end{aligned} \tag{2.44}$$

where D_{11} is a diagonal matrix and U and V are both orthonormal matrices. Note, k is the dimension of η_t and r is the number of nonzero singular values of $Q_2.\Pi$.

With all the information obtained so far, Proposition 1 of Lubik and Schorfheide (2003) shows that if a solution to (2.43) exists so that the endogenous forecasting errors η_t ($k \times 1$) can be expressed as a linear combination of the fundamental shocks ϵ_t ($l \times 1$) and the sunspot shocks ζ_t ($p \times 1$), it takes the following form

$$\eta_t = \left(-V_{.1} D_{11}^{-1} U'_{.1} Q_2.\Psi + V_{.2} M_1 \right) \epsilon_t + V_{.2} M_2 \zeta_t \tag{2.45}$$

where $V_{.2}$ is a $k \times (k - r)$ matrix, M_1 is a $(k - r) \times l$ matrix and M_2 is a $(k - r) \times p$ matrix. When $k = r$, the solution is unique and the second and the third terms of (2.45) drop out.

Parameter	β	G/Y	m/b	τ/b	G/b
Value	0.99	0.0838	0.2148	0.0754	0.0709

Table 2.1: Calibrated Parameter Values for Estimation

Name	Support	Density	Mean	St. Dev.	90% Interval
α_π	\mathbb{R}^+	Gamma	1.1	0.5	[0.43, 2.03]
γ_b	$(0, \beta^{-1})$	Uniform	0.51	0.29	[0.10, 0.91]
σ	\mathbb{R}^+	Gamma	2	1	[0.68, 3.88]
κ	\mathbb{R}^+	Gamma	2	1	[0.68, 3.88]
ν	\mathbb{R}^+	Gamma	2	1	[0.68, 3.88]
ϑ	$(0, 1)$	Beta	0.66	0.1	[0.49, 0.82]
ρ_Z	$(0, 1)$	Beta	0.7	0.1	[0.52, 0.85]
ρ_G	$(0, 1)$	Beta	0.7	0.1	[0.52, 0.85]
ρ_ϕ	$(0, 1)$	Beta	0.7	0.1	[0.52, 0.85]
ρ_ψ	$(0, 1)$	Beta	0.7	0.1	[0.52, 0.85]
σ_Z	\mathbb{R}^+	Inverse Gamma	1	1	[0.32, 2.45]
σ_G	\mathbb{R}^+	Inverse Gamma	1	1	[0.32, 2.45]
σ_ϕ	\mathbb{R}^+	Inverse Gamma	1	1	[0.32, 2.45]
σ_ψ	\mathbb{R}^+	Inverse Gamma	1	1	[0.32, 2.45]
σ_ζ	\mathbb{R}^+	Inverse Gamma	1	1	[0.32, 2.45]
M_1^Z	$(-3, 3)$	Uniform	0	1.73	[-2.40, 2.40]
M_1^G	$(-3, 3)$	Uniform	0	1.73	[-2.40, 2.40]
M_1^ϕ	$(-3, 3)$	Uniform	0	1.73	[-2.40, 2.40]
M_1^ψ	$(-3, 3)$	Uniform	0	1.73	[-2.40, 2.40]

Table 2.2: Prior Distributions for Model Parameters

Specification	$\ln p(Q_1^T \mathcal{M})$	Bayes Factor versus \mathcal{M}_1
\mathcal{M}_1 (AM/PF)	992	1
\mathcal{M}_2 (PM/AF)	847	$\exp(145)$
\mathcal{M}_3^a (PM/PF)	776	$\exp(216)$
\mathcal{M}_3^b (PM/PF)	910	$\exp(82)$

Table 2.3: Log Marginal Data Density Based on Pre-Volcker Period.

Note: \mathcal{M}_3^a : with a sunspot shock. \mathcal{M}_3^b : without sunspot shocks.

Specification	$\ln p(Q_1^T \mathcal{M})$	Bayes Factor versus \mathcal{M}_1
\mathcal{M}_1 (AM/PF)	801	1
\mathcal{M}_2 (PM/AF)	695	$\exp(106)$
\mathcal{M}_3^a (PM/PF)	589	$\exp(212)$
\mathcal{M}_3^b (PM/PF)	746	$\exp(55)$

Table 2.4: Log Marginal Data Density Based on Post-1982 Period.

Note: \mathcal{M}_3^a : with a sunspot shock. \mathcal{M}_3^b : without sunspot shocks.

	Pre-Volcker		Post-1982	
	Mean	90% Interval	Mean	90% Interval
α_π	9.8808	[8.0204, 11.9338]	9.5957	[7.7116, 11.6743]
γ_b	0.0593	[0.0178, 0.1346]	0.0576	[0.0172, 0.1284]
σ	10.6365	[8.0612, 13.5995]	12.9399	[9.8266, 16.4602]
κ	9.1293	[6.8070, 11.8668]	8.2644	[6.0064, 10.9363]
ν	1.3192	[0.8932, 1.8003]	0.9145	[0.4764, 1.3189]
ϑ	0.8331	[0.7764, 0.8816]	0.7508	[0.6649, 0.8255]
ρ_Z	0.9944	[0.9904, 0.9972]	0.9924	[0.9853, 0.9967]
ρ_G	0.8634	[0.8218, 0.9003]	0.8474	[0.7911, 0.9005]
ρ_ϕ	0.8417	[0.7899, 0.8868]	0.7567	[0.6751, 0.8282]
ρ_ψ	0.8901	[0.8234, 0.9455]	0.9091	[0.8570, 0.9524]
σ_Z	0.0617	[0.0488, 0.0770]	0.0558	[0.0441, 0.0700]
σ_G	0.1206	[0.0945, 0.1513]	0.1036	[0.0829, 0.1267]
σ_ϕ	0.0504	[0.0411, 0.0613]	0.0508	[0.0410, 0.0625]
σ_ψ	0.2229	[0.1876, 0.2638]	0.1530	[0.1242, 0.1862]

Table 2.5: Parameter Estimation Results.

Note: Posterior inferences for both periods are conditional on \mathcal{M}_1 (AM/PF).

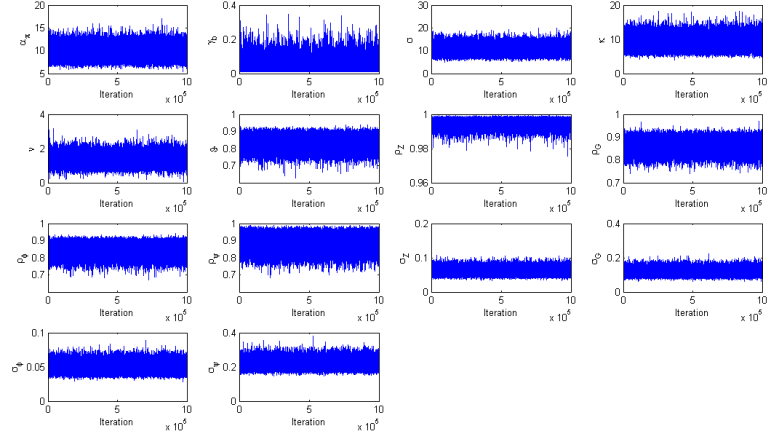


Figure 2.2: Trace Plots for \mathcal{M}_1 (AM/PF) Conditional on Pre-Volcker Period

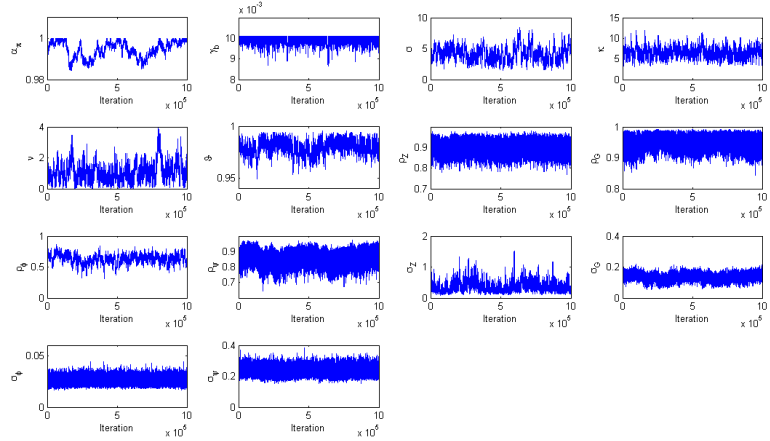


Figure 2.3: Trace Plots for \mathcal{M}_2 (PM/AF) Conditional on Pre-Volcker Period

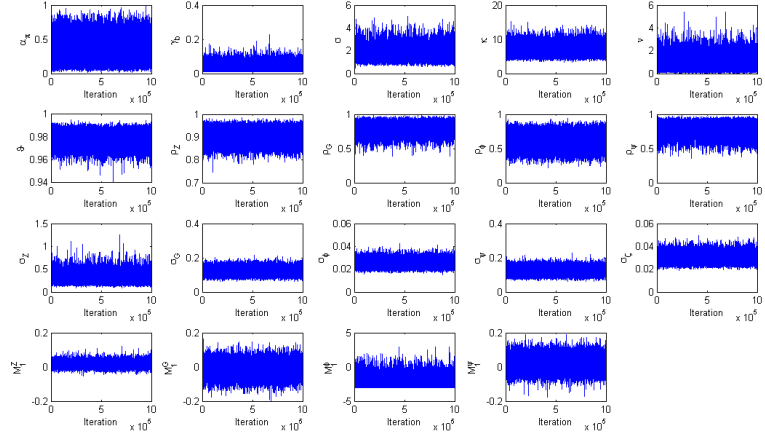


Figure 2.4: Trace Plots for \mathcal{M}_3 (PM/PF) Conditional on Pre-Volcker Period, With a Sunspot Shock.

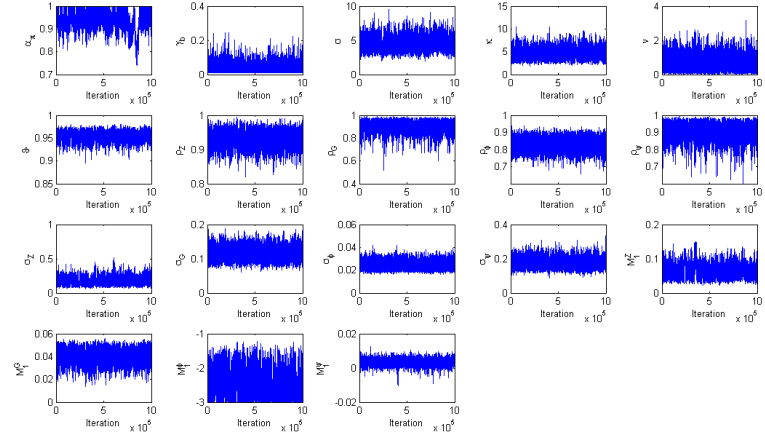


Figure 2.5: Trace Plots for \mathcal{M}_3 (PM/PF) Conditional on Pre-Volcker Period, No Sunspot Shocks.

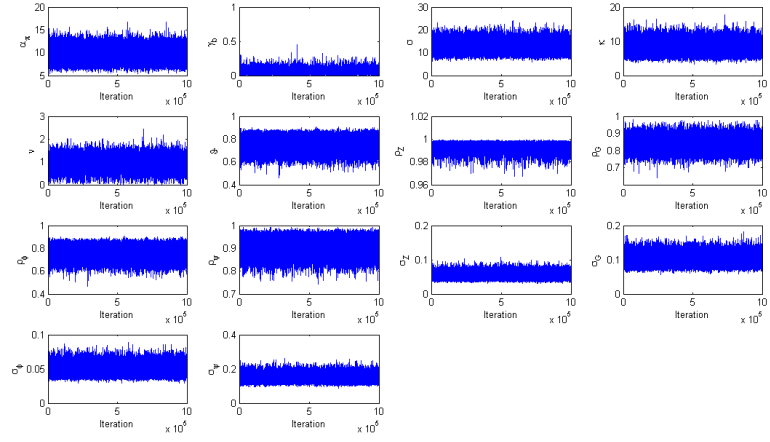


Figure 2.6: Trace Plots for \mathcal{M}_1 (AM/PF) Conditional on Post-1982 Period

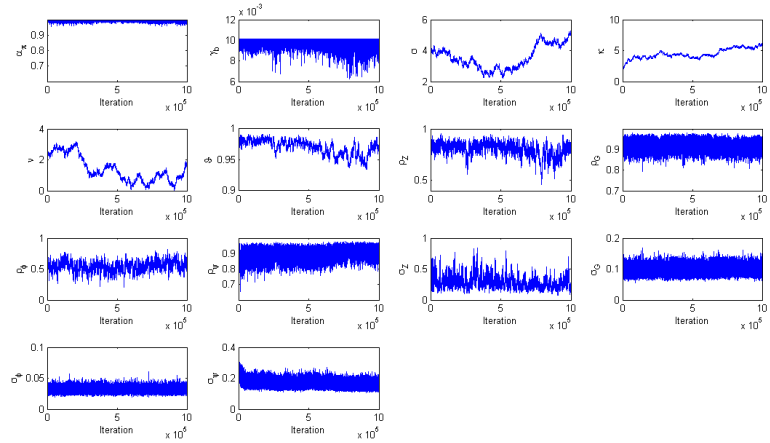


Figure 2.7: Trace Plots for \mathcal{M}_2 (PM/AF) Conditional on Post-1982 Period

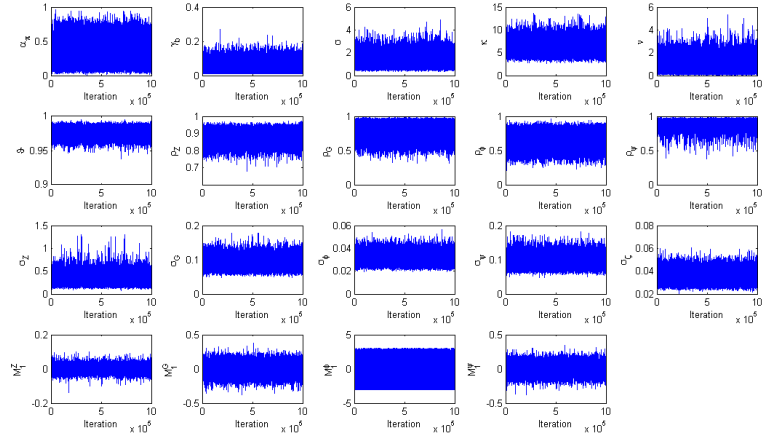


Figure 2.8: Trace Plots for \mathcal{M}_3 (PM/PF) Conditional on Post-1982 Period, With a Sunspot Shock

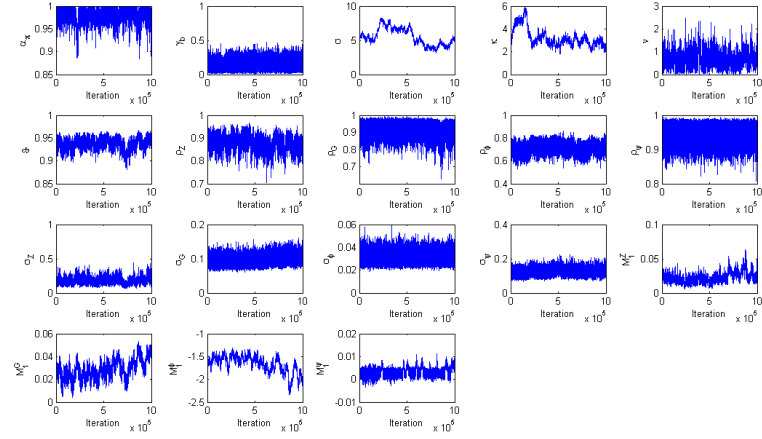


Figure 2.9: Trace Plots for \mathcal{M}_3 (PM/PF) Conditional on Post-1982 Period, No Sunspot Shocks

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Research Fields and Teaching Interests

- **Research:** Monetary Economics, Applied Macroeconometrics (especially estimation of DSGE models), Monetary and Fiscal Policy Interaction, Regime Switching in Macroeconomics, Open Economy Macroeconomics
- **Teaching (Undergraduate):** Principles and Intermediate Microeconomics, Principles and Intermediate Macroeconomics, Monetary Economics, Introductory Econometrics, Business Statistics, Mathematics and Quantitative Methods
- **Teaching (Graduate):** First-Year Macroeconomics, Monetary Economics, Open Economy Macroeconomics, Applied Macroeconometrics

Working Papers

- “On the Identification of Fiscal Policy Behavior,” CAEPR Working Paper No. 026-2008. (Job Market Paper)

Research In Progress

- “On the Identification of Fiscal Policy Behavior: A Bayesian DSGE Approach”
- “Bayesian Identification of Monetary and Fiscal Policy Behavior: A Regime-Switching DSGE Approach”
- “Investigating the Simultaneous Equations Approach in Identifying Fiscal Policy Behavior: A Simulation Study”

Conference and Seminar Presentations

- Econometric Society World Congress 2010, Shanghai Jiao Tong University, Shanghai, China, August 2010
- The 2009 Midwest Macroeconomics Meetings, Indiana University, Bloomington, IN, May 2009
- The 9th Annual Missouri Economics Conference, University of Missouri-Columbia, Columbia, MO, March 2009
- The 2009 Midwest Economics Association Annual Meeting, Cleveland, OH, March 2009
- The Third Annual Economics Graduate Student Conference, Washington University in St. Louis, St. Louis, MO, October 2008
- Jordan River Conference, Indiana University, Bloomington, IN, April 2007
- Macro Group Brown Bag Seminar organized by Eric Leeper, Indiana University, Bloomington, IN, February and October 2007, January and March 2008

Teaching Experience

- **ECON 370 Statistical Analysis for Business and Economics:**
 - ◇ Associate Instructor (AI with Full Teaching Responsibility): Summer I 2007, Spring 2008, Summer I 2008, Summer I 2009
 - ◇ Associate Instructor (AI with Full Teaching Responsibility for Weekly Lab Sections): Fall 2005, Spring 2006, Fall 2006, Spring 2007, Fall 2007, Fall 2008, Fall 2009
 - ◇ Graduate Assistant (GA): Fall 2004, Spring 2005
 - ◇ Presenting at Pre-Semester New AI Training as Experienced AI: August 2006, January 2007, August 2008
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